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1 Introduction

Since more than 50% of jobs are informal and pay no social security in the developing world, in the last 20 years governments in 60 countries worldwide paid up to 1% of GDP to enact social pensions with the aim of mitigating old age poverty (Bosch *et al.*, 2013). In this paper we try to gain a better understanding of the reasons why a social pension enacted in 2008 in Bolivia, *Renta Dignidad*, has a take-up of only about 60%, which poses a challenge to achieve its objective of mitigating poverty by paying about 25% of per capita income or 60 US dollars in PPP (Bosch *et al.*, 2013). First, we test whether the gender of eligible elderly spouses affects take-up within a household, to better understand whether gender differences in spouses' ability to take all necessary administrative steps to obtain the pension or complementarities in the event of eligibility by both spouses are at play. This is instrumental to then study whether eligibility leads to a decrease in poverty incidence, which is about 50% when considering households with elderly spouses in Bolivia.

Second, we assess whether eligibility induces behavioural responses and whether they are gendered if spouses' preferences exhibit high enough heterogeneity, which may lead to a bargaining process between spouses, or whether they exhibit gender complementarities when both spouses are eligible and may coordinate on how to use the pension income. We focus on consumption and labour supply, i.e. direct responses, because the pension income shock may allow to buy the same consumption bundle with a lower labour supply, to keep labour supply unchanged and afford more consumption, or a combination of both. In addition, the pension income may induce indirect behavioural responses. Since extended families are a frequently observed arrangement to cope with children and elderly care in developing countries, we focus on testing whether eligibility modifies the composition of a household.

We identify the effect separately by spouse's gender thanks to a bi-dimensional regression discontinuity design (RDD), with spouses' age as forcing variables, since individuals become eligible when turning 60. We estimate the effect by using data from the 2008 and 2009 waves of the Bolivian household survey. We find that eligibility leads to a significant "jump" in take-up (from 0% to about 60%). The only significant gender differences are observed when we consider heterogeneity by wealth, with take-up being lower for eligible females while higher when both spouses are eligible in wealthy households. However, the observed differences in take-up are not sizable enough to lead to a proportional decrease in poverty incidence. Similarly, when we look at consumption and labour supply we do not find a significant eligibility effect by spouses' gender.

When we look at household composition, we find that the probability of observing co-residing grandchildren is higher when both spouses are eligible (50 pp or 280%), while that of observing an elderly couple's adult children is unchanged. Importantly for our research design validity, the distribution of spouses' age and baseline characteristics are continuous at the 60 cutoff, which rules out sizable sorting by individuals in or out of the pension. Additional tests show that our results are robust to using a different age range than 50-70 and are not confounded by an anticipation effect, i.e. borrowing before becoming eligible to then repay using the pension income, or by a placebo effect, i.e. behavioural responses by spouses turning 60 before the pension was enacted.

Our paper contributes to the growing number of related studies on social pensions in Mexico and in Latin America and, in particular, to two closely related studies on Bolivia (Escobar Loza *et al.*, 2013; Hernani-Limarino and Mena, 2015). Most of them find, although using different empirical specifications, that consumption increases while labour supply decreases for eligible individuals. Our paper complements them by focusing on eligibility by spouses' gender in a household and by finding evidence of complementarity in spouses' eligibility effect on a decision over household composition, namely co-residing grandchildren.

In addition, we contribute to two studies that have identified gender effects using data on the South African Old Age Pension (Edmonds, 2006; Ardington *et al.*, 2009) and find evidence of an increase in children's schooling and in young adults' employment, respectively. Our paper complements them by showing that in a poorer country than South Africa a less generous social pension induces a qualitatively similar indirect effect on non-beneficiaries although more muted as eligibility leads only to an increase in time invested by grandparents in their grandchildren's care.

The structure of the rest of the paper is as follows. Section 2 puts our paper in context in the related literature. Section 3 describes the institutional setting and the data. Section 4 describes the empirical research design and its validity while section 5 presents the results. Finally, section 6 discusses the results and concludes. Additional results can be found in the Appendix.

2 Literature review

Our paper complements related studies on social pensions in Mexico and in a number of countries in Latin America including Bolivia as, differently from all of them, it explicitly looks at the role played by spouses' gender and household wealth in explaining differences in take-up and decisions over the amount and type of expenditure induced by eligibility for the Renta Dignidad pension. Social pensions enacted in these regions over the last 20 years have eligibility rules based on age cutoffs, they were discontinued over time in Bolivia and they are restricted to residents of a region in Mexico.

These rules make estimating the eligibility effect for one or more household members possible by using either regression discontinuity, difference in difference, propensity score matching or a combination of these designs. This is typically done by focusing on the age of the eldest member in a household to compare food expenditure in households in which the eldest member's age is past the eligibility cutoff with others in which the eldest member is younger, accounting for incomplete take-up where relevant and focusing only on households whose eldest member age is close to the cutoff in the case of regression discontinuity. In our paper we, instead, compare the behaviour of the following groups of households: those in which both spouses are eligible for a pension, those in which only one spouse is eligible, with particular attention to the role played by gender of the eligible spouse, and those in which no spouse is.

Our paper is closely related to two studies on the Renta Dignidad pension. Escobar Loza *et al.* (2013) study the eligibility effect by using a regression discontinuity design in which the forcing variable is the age of the eldest member of a household. The data used was obtained on purpose for an official evaluation of the pension by using the pool of households from the household survey and oversampling those with the eldest member in the age range 55-65. The main results are that being eligible reduces poverty and, in addition, increases household consumption. Hernani-Limarino and Mena (2015) is a related study that uses a refinement of a difference in difference design, called change-in-change, to identify the counterfactual outcome for non-eligibles and study quantile treatment effects. This is done by exploiting the change in cutoff value from 65 to 60 when Renta Dignidad replaced in 2008 Bonosol, a similar social pension whose age eligibility cutoff was 65. Their main findings are a positive effect of non-labor income and, for women, a negative one on their labor supply.

Our paper is also related to studies quantifying the effect of Bonosol, the social pension that was replaced by Renta Dignidad in 2008. Martinez (2004) studies the effect of receiving at least one pension in a household in the period 1998-2002. This is done by comparing outcomes of interest for households in a period in which the pension was discontinued, 1998-1999, and in a subsequent period in which payments were resumed, 2000-2002, by way of a difference in difference design used in combination with a regression discontinuity design to compare (in)eligible households close to the cutoff. The main results show that consumption is higher, particularly in rural areas, and, in addition, school enrollment for children is higher. Yanez-Pagans (2008) studies the effect of the same pension on children's schooling by using data on 2001 and a regression discontinuity design to compare (in)eligibles. The main results are that children's schooling expenditure is higher, particularly for non-indigenous women.

In addition, our paper is related to studies of the (in)direct effect of non-contributory pensions in Mexico and in Latin America. The impact of the Adultos Mayores pension in Mexico, whose eligibility rules are being 70 or older and residing in the Federal District in Mexico City, has been studied by Salinas-Rodríguez *et al.* (2014) who use a difference in difference design to exploit the residential eligibility criterion and find that eligibility decreases depression and increases subjective views on empowerment. Galiani *et al.* (2016) combine a difference in difference with a regression discontinuity design to exploit both eligibility criteria and find that mental health improves, a switch is observed from paid work to work in family businesses and consumption is higher. Gutierrez *et al.* (2017) focus on school enrollment for beneficiaries' co-residing grandchildren, use a regression discontinuity design to exploit the age cutoff and find a positive effect.

When we look at other countries, Pension 65 is a social pension enacted in Peru for individuals who are 65 or older and live in a household below the poverty line. Its effect has been studied by Bando *et al.* (2016) who use a regression discontinuity design to exploit the poverty-based eligibility criterion and find that eligibility decreased depression and labour supply while it increased consumption. In a related study Novella and Olivera (2017) find with linear and non-linear regressions a negative association between the score in a cognitive test and retirement. Although the retirement measure does not distinguish different pension types, most elderly in Latin America only receive a social pension. In Brazil, the Benefício de Prestação Continuada pension was enacted for individuals who are 65 or older. de Oliveira *et al.* (2017) estimate its effect by using a regression discontinuity design to compare (in)eligibles and find a decrease in the labor force participation of the elders.

Finally, our paper is related to two studies assessing the gender-specific impact of the Old Age Pension enacted in South Africa in the 1990s for individuals who are 60 or older. Edmonds (2006) studies the eligibility effect on school age children by looking at whether the eldest man or woman are eligible by way of a regression discontinuity whose two forcing variables are the age of the eldest female and male in a household. The main results show a large increase in schooling attendance and a decline in total hours worked when black South African families are eligible for the pension, with the schooling effect being more pronounced when a male is eligible. In a related study Ardington *et al.* (2009) quantify the pension effect on labor supply of prime-aged adults by using longitudinal data to compare households becoming eligible and those no longer eligible due to changes in the household composition over time. They find that the pension increases employment among young adults, which occurs primarily through labor migration. In addition, while young adults migrate independently of gender when a woman is eligible, only young male adults migrate when a man is eligible. This evidence suggests that the pension relaxed household liquidity constraints and that preferences may be heterogeneous by gender.[1]

3 Institutional setting and data

Section 3.1 describes the institutional setting of the pension system in Bolivia, to clarify how the pension works, while Section 3.2 describes the data used in the empirical analysis.

3.1 Pension system in Bolivia

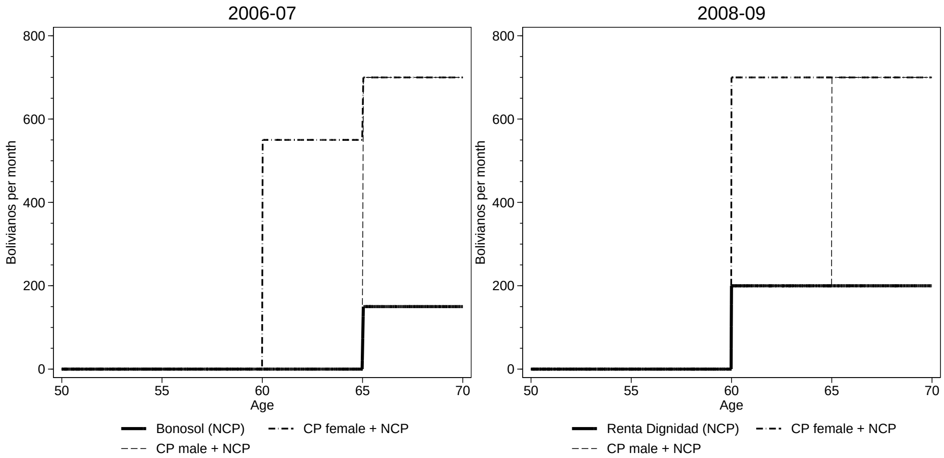
Renta Dignidad is not the first non-contributory pension in Bolivia. In 1997, the government enacted Bonosol, a pension paying all citizens who were 65 or older 1,300 bolivianos per year, that was about 27% of per capita income and 85% of income for those living in extreme poverty (von Gersdorff, 1997). Bonosol was part of a broader social and economic reform agenda with two main aims. The first was reducing high income inequality in the country, that was in the top quartile of the distribution of countries worldwide measured using the GINI index (CIA World Factbook, 2014). The second was dealing with the consequences of the high share of informal employment, about 60% (World Bank, 2009), which is detrimental as it does not lead to the accumulation of contributory pension rights over time. Bonosol was not paid in the period 1998-2000, as it was judged financially untenable. In 2001, instead, the pension was resumed although its amount decreased to about 420 bolivianos in 2001-2002, while it increased to 1,800 bolivianos in 2003. Finally, the pension was discontinued in December 2007.[2]

[1] Duflo (2003) studies the impact of the South African pension by comparing households with no pension with those in which a female received it and, also, with those in which a male received it by instrumenting pension receipt with eligibility. The main results are a positive impact on anthropometric status, i.e. weight for height and height for age, for girls but little effect for boys, although only when the pension recipient is a woman. The main limitation in Duflo (2003)'s identification of gender effects is not assessing whether results are robust to considering only individuals close to the eligibility age cutoff. Additional studies on the same pension focus on its impact on employment or on living arrangements, although either only partly or not at all accounting for the role of beneficiaries' gender (Bertrand *et al.*, 2003; Edmonds *et al.*, 2005; Posel *et al.*, 2006; Hamoudi and Thomas, 2014; Tondini *et al.*, 2017).

[2] See Willmore (2006) for additional information about pension reforms in Bolivia.

Renta Dignidad was enacted on 1st February 2008 with two main differences with respect to Bonosol. First, the age eligibility cutoff decreased to 60. Second, the amount paid increased to 2,400 bolivianos per year, except for individuals obtaining a contributory pension and public sector employees (about 20% of all beneficiaries), who received 1,800 bolivianos.[3]

Figure 1: Pension types by age eligibility



In addition to non-contributory pensions, the Bolivian government pays contributory pensions to individuals who paid social security throughout their working life. Retirement age for females was 60 while it is 65 for males. However, it turns out that only about 40% of workers are in the formal sector and about 20% obtain a contributory pension in the age range 60-65.

Figure 1 shows the different sources of pension income that an individual obtained per month, measured in bolivianos on the vertical axis, as a function of age on the horizontal axis. The left-hand side panel shows pension income for the period 2006-2007, when the Bonosol non-contributory pension, abbreviated NCP in the figure, was in place. Individuals younger than 60 were not eligible for any pension income. When turning 60 females

[3]The first pension payment was made up to a month after an individual turned 60, at either a bank or an authorized military enclosure subject to identity verification. Alternatively, arrangements were in place to obtain it at home. The pension was paid on a monthly basis, except for individuals obtaining 2,400 bolivianos who could choose either monthly or less frequent payments. Additional information is available in Escobar Loza *et al.* (2013).

who had paid social security throughout their working life were eligible for a contributory pension, abbreviated CP in the figure, of 550 bolivianos per month while males were only eligible when turning 65. In addition, when turning 65, both females and males were eligible for Bonosol, which consisted of 150 bolivianos per month. The right-hand side panel shows the same information although for the period 2008-2009, when Bonosol was replaced by Renta Dignidad. Similarly to 2006-2007, individuals younger than 60 were not eligible for any pension income. At 60, all individuals became eligible for Renta Dignidad, paying 200 bolivianos per month, and females who had paid social security also received a contributory pension. Finally, at 65, males who had paid social security also received a contributory pension.

3.2 Data

We use data from the 2008 and 2009 waves of the Bolivian household survey (Encuesta de Hogares) run by the Bolivian National Statistics Institute. Our unit of observation in the data is a household since our treatment of interest is whether one or more spouses were eligible for the Renta Dignidad pension in a household and (in)eligibles' gender. Spouses are defined by using data on who is the head of the household and who is her/his spouse, on their gender and on their marital status.[4]

Testing our hypothesis that the effect of pension eligibility differs by spouses' gender requires variation in spouses' age across households around the 60 age eligibility cutoff. Figure 2 measures for each household, shown as circles, the female spouse's age on the vertical axis and the male's on the horizontal one. Although the figure shows that households with either no spouse or two spouses eligible for a pension are relatively more frequent, respectively 25% and 45%, the share of households with a spouse eligible is non-negligible, with about (10)20% in which only the (fe)male spouse is eligible.

Following Escobar Loza *et al.* (2013), we define as *eligible* for a pension a spouse in a household if (s)he was 60 years and a month old since up to a month elapsed between the time an individual turned 60, submitted the application for the Renta Dignidad pension and the time the first pension

[4]The data can be downloaded from section "Banco de Datos" in `www.ine.gob.bo`. Alternatively, a copy of our dataset is available upon request.

Figure 2: Variation in spouses' age in a household

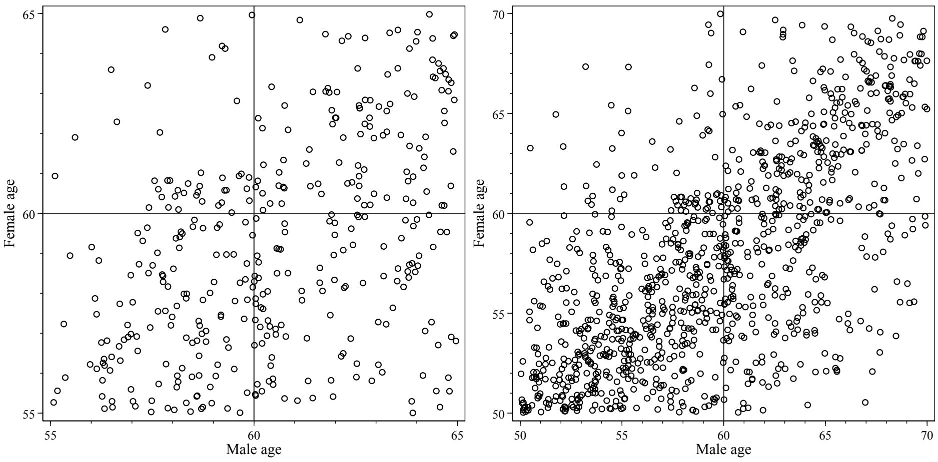
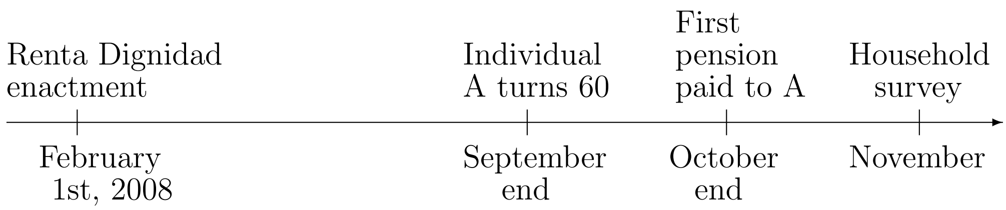


Figure 3: Measurement of pension eligibility and outcomes in 2008



payment was made, as illustrated in Figure 3. In addition, in a setting in which eligibility is based on age and our outcomes of interest are measured in the yearly household survey in early November, eligible spouses are those who were 60 or older on 30th September, i.e. one month before our outcomes of interest were measured in the survey. Non-eligibles are, instead, those who turned 60 later. Since the survey contains information on day, month and year of birth, we can create a precise measure of age in which the integer part measures age in years while the decimal one measures fractions of years.[5]

[5]Eligible spouses in the 2008 (2009) wave of the survey are those who were born on 30th September 1948 (1949) or earlier.

We selected the data sample used in our empirical analysis as follows. We start by considering the full set of 30,695 individuals in the 2008 and 2009 waves of the Bolivian household survey. Subsequently, we collapse the data at the household level, obtaining 7,974 households. Since in our analysis we focus on households whose spouses have an age close to the 60 cutoff determining eligibility for Renta Dignidad pension, we restrict our attention to the 1,048 households in which the eldest individuals of both sexes are in the 50-70 age range. Finally, when we further restrict our focus only on those households with two spouses who are both in the 50-70 age range, the number of observations is unchanged, thus suggesting that all houses in which the eldest female and male are in the 50-70 age range are households with a married couple with both spouses in that age range. We chose not to restrict the age range further in our empirical analysis to ensure that our main results do not suffer from misspecification due to a number of observations too low to estimate an empirical specification that allows for heterogeneity by spouse gender in the pension eligibility effect.

We use two *poverty* measures defined by the National Statistics Institute in Bolivia. The first is a dummy equal to 1 if total income per capita is below the poverty threshold, set at 418 bolivianos in the survey data for 2008 and 2009. The second one is a dummy equal to 1 if total income per capita is below the extreme poverty threshold, set at 218 bolivianos over the same period. We measure spouses' *labour supply* thanks to a dummy set equal to 1 in the survey data if a given spouse worked over the week before the survey interview and 0 otherwise. In addition, we define a dummy to capture whether at least one spouse worked.[6]

We also include two typically important sources of expenditure for a household, measured in bolivianos over the month before the survey was held. The first is total food expenditure, which we obtained by summing expenditure over all food items, from raw to pre-cooked food. The second is educational expenditure, which we computed to assess whether households with one or more spouses eligible for a pension contribute to finance their grandchildren's education more than households with no eligible spouse. Both measures of expenditure are expressed in bolivianos and refer to the

[6]Poverty threshold values were defined based on the cost of basic food and non-food consumption needs. Additional details about the survey design can be found in www.ine.gob.bo.

month before the survey interview. We dropped 3 observations whose value of food consumption was about 100,000 bolivianos as it is suspiciously high and most likely due to coding error. Finally, we obtained measures of household composition by setting a dummy equal to 1 if adult sons or daughters co-reside with their elderly parents and a dummy equal to 1 if grandchildren lived in the household.

The predetermined characteristics used in our empirical analysis are, firstly, a dummy for whether one or more spouses in a household belongs to the Quechua minority ethnicity and a dummy created analogously for the Aymara ethnicity. In addition, we created dummies for whether at least a spouse had completed 5 years of education (primary), 8 years (secondary) or 12 years (post-secondary). To control for a predetermined measure of health we also created a dummy for whether at least one spouse in a household was ill in the 4 weeks before the yearly survey was administered in a household. This measure is predetermined since *i*) we define as eligible for a pension those spouses turning 60 at the end of September in a year and *ii*) it took approximately a month between the 60th birthday and the first pension payment and the household survey was conducted between the end of October and the beginning of November. In other words, the health dummy is equal to 1 if a spouse was ill between the end of September and the end of October. Note that a spouse may be ill after turning 60 but, crucially, before being eligible for a pension. We also created a measure of household wealth by way of a dummy equal to 1 if at least one spouse had a private health insurance, which is typically granted by stable jobs with above median remuneration. Finally, we created a dummy to distinguish households in urban areas from those in rural ones and a dummy for those households who participated in the survey in 2009.[7]

Table 1 shows summary statistics of our outcomes of interest and of baseline characteristics for households with spouses in a 50-70 age range. This is done separately by whether no spouse is eligible for a pension when they are all younger than 60, only one of the two spouses is 60 or older and therefore eligible, distinguishing by gender, and, finally, all are 60 or older

[7]Since a free health insurance (Seguro de Salud para el Adulto Mayor) was offered since 2006 as a universal program to all individuals who were 60 or older, for these individuals our health insurance dummy is set equal to 0 as they did not pay for health insurance.

and therefore eligible.

Table 1: Summary statistics

	Both 50-59	Females 60-70, Males 50-59	Females 50-59, Males 60-70	Both 60-70
		Outcomes		
Take-up female	0.02	0.76	0.05	0.81
Take-up male	0.01	0.02	0.76	0.84
Take-up amount (bol.)	6.45	155.39	153.99	313.11
Poverty	0.45	0.51	0.48	0.47
Extreme poverty	0.28	0.25	0.27	0.22
Female working	0.64	0.45	0.59	0.47
Male working	0.91	0.88	0.82	0.79
1+ working	0.95	0.89	0.90	0.84
Food exp. (bol.)	276.78	290.22	301.49	352.42
Food exp. (%)	0.37	0.47	0.42	0.49
Educ. exp. (bol.)	209.76	259.77	143.22	80.98
Educ. exp. (%)	0.15	0.12	0.10	0.06
Co-residing adult children	0.82	0.57	0.66	0.48
Co-residing grandchildren	0.17	0.34	0.30	0.36
		Predetermined characteristics		
1+ Quechua ethnicity	0.35	0.30	0.38	0.41
1+ Aymara ethnicity	0.33	0.42	0.33	0.30
1+ educ. 5+ years	0.66	0.64	0.53	0.49
1+ educ. 8+ years	0.45	0.39	0.32	0.29
1+ educ. 12+ years	0.29	0.27	0.26	0.21
1+ ill	0.43	0.54	0.57	0.55
1+ health ins.	0.31	0.33	0.33	0.27
1+ contrib. pension	0.06	0.14	0.17	0.23
Urban	0.61	0.59	0.52	0.45
Year 2009 survey	0.49	0.48	0.54	0.49
N	498	67	210	273

The top panel in Table 1 shows that Renta Dignidad take-up increases discontinuously when at least one spouse is eligible, with non-compliance being small at about 2%. It also shows that the incidence of poverty is about 45% when no spouse is eligible, with a slight increase when at least one spouse is eligible. The incidence of extreme poverty is lower than 30%, being 28% when no spouse is eligible and lower, at 22% when all are eligible. When we look at labour supply, it is higher for males over females, ranging from 91% and 64% respectively when no spouse is eligible to 79% and 47% when all are eligible. When we look at food expenditure, it is approximately 277 bolivianos or 37% of total expenditure when no spouse is eligible and it increases to 352 or 49% when both spouses are eligible. Finally, educational expenditure ranges from about 210 bolivianos or 15%

when no spouse is eligible to 81 or 6% when all are eligible.

When we look at household composition, the probability of observing young adults co-residing with their parents is 82% when no spouse is eligible, while it is substantially lower when at least one spouse is eligible. When we look at the probability that grandchildren co-reside with their grandparents, it is 17% when no spouse is eligible and it is higher otherwise, with a substantial increase to 36% when both spouses are eligible.

The bottom panel in Table 1 shows means of baseline characteristics. In 30-40% of households at least one spouse is of Quechua ethnicity with no sizable difference between the groups defined by spouse eligibility and similar values hold for Aymara ethnicity. At least one spouse has completed at least compulsory education (5 years) in 49-66% of households, at least one has completed at least secondary (8 years) in 29-45% of households, and at least one spouse completed post-secondary education in in 21-29% of households.

In addition, the bottom panel in Table 1 shows that the probability of a sick spell for at least one spouse ranges from 43% for households with no spouse eligible to 57% for households in which only males are eligible. As for the frequency of households paying for a health insurance out of their pocket or through their employer, which is either public (Caja de Salud) or private, it ranges from 27% when both spouses are eligible to 33% when only a spouse is. We also report the frequency of households with at least one spouse receiving a contributory pension, which ranges from 6% when no spouse is eligible to 23% when all spouses are eligible and the frequency of urban households, which varies from 45% when all spouses are eligible to 61% when no spouse is eligible. Finally, the frequency of households in the data from the 2009 wave of the survey is close to 50% for all the groups we consider by spouses' eligibility.[8]

4 Research design and validity

We describe the empirical research design that we use to estimate whether the positive income shock induced by spouses' eligibility for the Renta

[8]We discuss further evidence on differences in household baseline characteristics by pension eligibility status in section 4.2.

Dignidad pension and whether gender differences play a relevant role in influencing household decisions in section 4.1. In addition, we discuss our research design validity and offer evidence in support of it in section 4.2.

4.1 Regression discontinuity design

A linear regression of an outcome Y on a dummy D equal to 1 if an individual is eligible for a pension and 0 otherwise leads to spurious estimates if individuals' unobservable characteristics correlate with eligibility status. By exploiting the age-based cutoff to be eligible for the pension we can, instead, estimate a clean effect of pension eligibility locally at the age 60 cutoff thanks to a regression discontinuity design (RDD) since, by focusing on those individuals whose age is close enough to 60, eligibles and ineligibles are fully comparable in their predetermined characteristics. Since not all those who are eligible for the pension apply for it, for a variety of reasons, take-up may be endogenous.

We follow the specification used by Edmonds (2006) to study eligibility for the South Africa Old Age Pension in two important dimensions. The first is focusing on eligibility for the pension and identify the intention to treat effect (ITT) at the cutoff since take-up is endogenous. With incomplete take-up at the age 60 cutoff, the ITT effect is a lower bound of the pension effect on those who actually receive it, the local average treatment effect (LATE). The second is focusing on eligibility separately for each spouse as as decisions over consumption, labour supply or household composition are typically the outcome of a negotiation between spouses and may be influenced by whether no spouse, one or both of them are eligible for a pension, with potential gender differences.

Equation (1) shows the empirical specification that we use to separately estimate the pension eligibility effect by spouse's gender on our outcome of interest Y. We let female spouse's age be measured by A_F, with the integer part measuring age in years while the decimal one fractions of years. Eligibility is measured by a dummy $D_F = \{A_F \geq 60\}$ equal to 1 if she is 60 or older, with age and the eligibility dummy being created analogously for male spouses. Hence, β_1 and β_2 in equation (1) capture the eligibility effect when only one spouse is eligible (female and male, respectively) while β_3 captures the difference in the effect when both spouses are eligible for the pension.

$$
\begin{aligned}
Y &= \beta_0 + \beta_1 D_F + \beta_2 D_M + \beta_3 D_{FM} + \\
&\quad + f(A_F - 60, A_M - 60, D_F, D_M) + U
\end{aligned} \tag{1}
$$

The polynomial $f(\cdot)$ in equation (1) accounts for spouses' age trends in our outcomes of interest and is defined in equation (2), with age rescaled at 60. Parameters from β_4 to β_8 capture own age trends and are allowed to differ by whether own age is 60 or greater while $\beta_9 - \beta_{15}$ capture trends in spouse's age or in spouses' joint age.

$$
\begin{aligned}
f(\cdot) &= \beta_4(A_F - 60) + \beta_5(A_M - 60) + \beta_6(A_F - 60)(A_M - 60) + \\
&\quad + \beta_7 D_F(A_F - 60) + \beta_8 D_M(A_M - 60) + \beta_9 D_M(A_F - 60) + \\
&\quad + \beta_{10} D_F(A_M - 60) + \beta_{11} D_F(A_F - 60)(A_M - 60) + \\
&\quad + \beta_{12} D_M(A_F - 60)(A_M - 60) + \beta_{13} D_{A_F} D_{A_M}(A_M - 60) \\
&\quad + \beta_{14} D_{A_F} D_{A_M}(A_F - 60) + \beta_{15} D_{A_F} D_{A_M}(A_M - 60)(A_F - 60)
\end{aligned} \tag{2}
$$

Our research design is a bi-dimensional RDD as we use two running variables, female and male spouses' age, to assess the role of eligibility by spouse's gender in influencing our outcomes of interest. Edmonds (2006) is the only other study that uses a similar empirical methodology, to the best of our knowledge, although it differs from ours as the two running variables in Edmonds (2006) are age of the eldest man and woman in a household.

The RDD *identifying assumption* is absence of or imperfect sorting by individuals on either side of the age 60 cutoff. Although the identifying assumption is untestable, date of birth and in our setting whether an individual is barely younger or older than age 60 cutoff on 30th September 2008, the date we use to define eligibles in 2008, are arguably exogenous. In section 4.2, we carefully assess whether the distribution of spouses' age, as well as other predetermined characteristics are continuous at the 60 cutoff, finding evidence in support of it.

Finally, we assess whether take-up and the pension eligibility effect vary by household wealth as evidence in the literature suggests that poor households tend to be more credit constrained than others (e.g. Martinez, 2004; Edmonds, 2006). We define as a proxy for household wealth a dummy equal to 1 if one or more spouses in a household pays for a health insurance policy which, as we discussed in section 3, is more likely to be affordable by households with stable jobs or by affluent ones.[9]

We estimate equation (1) using a local linear regression in spouses' age rescaled at the 60 cutoff. In all specifications we use a rectangular kernel of size 10 years, i.e. we estimate a linear regression using a data sample obtained by considering those households in which both spouses' age is within 10 years on either side of the age 60 cutoff. This kernel has been chosen due to the low number of observations to estimate our empirical design in the survey data and, in addition, to its simplicity with respect to more sophisticated ones, since kernel choice tends to have little impact in practice (Imbens and Lemieux, 2008). In addition, the standard error formula used is robust to heteroskedasticity and we also correct standard errors by using the sampling weights in the survey. Finally, we add as baseline characteristics dummies for ethnicity, education, health status and insurance, eligibility for a contributory pension, residential area and survey year.[10]

4.2 Research design validity

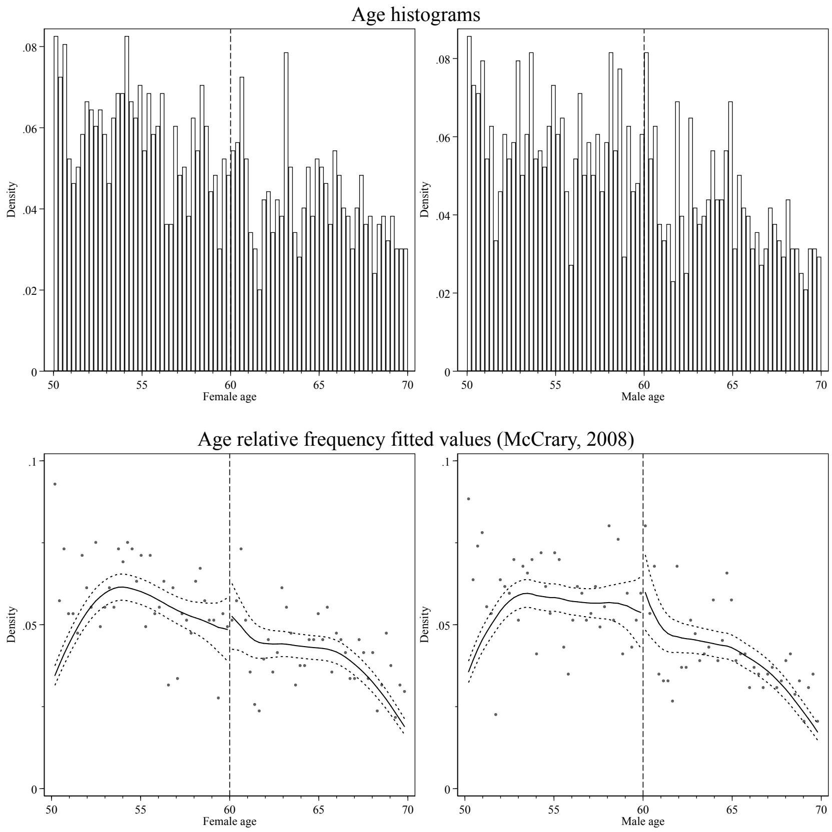
The untestable RDD *identifying assumption* is that individuals are unable to sort themselves into the treated group, or out of it, by manipulating, for example, their age or date of birth. We offer evidence in support of this assumption by assessing empirically whether the age distribution by spouse gender is smooth at the age 60 cutoff and, similarly, whether individuals' baseline characteristics are balanced at the cutoff.[11]

[9] We estimate the heterogeneous pension effect by wealth by letting D_{HI} be a dummy equal to 1 if at least one spouse in a household has a health insurance and by multiplying the terms in equation (1) by $(1 + D_{HI})$. The parameters associated to $D_F \cdot D_{HI}$, $D_M \cdot D_{HI}$ and $D_{FM} \cdot D_{HI}$ measure the difference in the pension eligibility effect for wealthy households relative to others.

[10] Sample weights are the same when data are used at the individual or at the household level since the number of households in the primary sampling unit in the survey varies from 80 to 350 based on population density.

[11] See Lee and Lemieux (2010) for a discussion of RDD identifying assumption and validity.

Figure 4: Age distribution



The top panel in Figure 4 shows age histograms, separately by spouse gender in a household, with the histogram bin set to 90 days to ensure that discontinuities can be detected visually. Visual inspection suggests no suspicious "jump" in histogram bins height at the cutoff, hence supporting the validity of the research design. This result is confirmed by density-based tests of the null hypothesis of no manipulation (McCrary, 2008). They are reported in the bottom panel in Figure 4 and show that confidence intervals of the difference in age density fitted values, indicated as dashed lines in the figure, overlap at the 60 cutoff. While the main diagnostic for sorting of individuals at the cutoff in a RDD excludes significant sorting, we observe a number of spikes in the histograms in the top panel in Figure 4 at a number of integer age values such as 59 and 61. Although clearly identifying the determinants of such spikes is beyond the scope of our analysis, we speculate that they may be due to demographic factors, since we measure age at the end of September 2008 and 2009.

Table 2: Balance of baseline characteristics at the age 60 cutoff

	1+ ethnicity		Education (years)		
	Quechua	Aymara	5+	8+	12+
D_F	0.156	-0.137	0.290	0.085	0.140
	(0.172)	(0.181)	(0.178)	(0.180)	(0.166)
D_M	0.002	-0.154	0.267*	0.148	0.041
	(0.151)	(0.149)	(0.161)	(0.144)	(0.130)
D_M*D_F	-0.053	0.026	-0.535**	-0.230	-0.228
	(0.242)	(0.238)	(0.245)	(0.231)	(0.211)
N	1,048	1,048	1,048	1,048	1,048
Bandwidth	10	10	10	10	10
$JE = D_F + D_M + D_F * D_M$	0.105	-0.266	0.022	0.003	-0.047
P-value (JE=0)	0.538	0.085	0.894	0.980	0.706
Mean value (incl. HH)	0.333	0.321	0.661	0.452	0.315

	Health		Contributory	Household characteristics	
	sick spell	insurance	pension	urban	survey in 2009
D_F	-0.143	-0.027	0.066	-0.298	-0.003
	(0.176)	(0.164)	(0.138)	(0.185)	(0.183)
D_M	-0.125	0.056	0.128	-0.222	0.013
	(0.154)	(0.142)	(0.100)	(0.157)	(0.160)
$D_M * D_F$	0.020	-0.014	-0.135	0.105	-0.210
	(0.239)	(0.216)	(0.180)	(0.246)	(0.246)
N	1,048	1,048	1,048	1,048	1,048
Bandwidth	10	10	10	10	10
$JE = D_F + D_M + D_F * D_M$	-0.248	0.015	0.058	-0.415	-0.201
P-value (JE=0)	0.126	0.914	0.556	0.011	0.218
Mean value (incl. HH)	0.426	0.299	0.058	0.476	0.574

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

In the absence of manipulation we also expect no difference in baseline characteristics when we compare individuals who were barely younger than 60 with those who were barely older, as such characteristics are predetermined. Table 2 shows estimates obtained by regressing each of our baseline characteristics, whose list is reported in Table 1, on age dummies and a RDD polynomial in age using equation (1). For each baseline characteristic, we report estimates of the eligibility dummies by spouse gender (D_F and D_M), as well as the joint eligibility effect ($D_F + D_M + D_F * D_M$), together with the p-value of the null hypothesis of no joint effect. Overall, Table 2 shows that differences in baseline characteristics at the cutoff are small and not significant. The same results are obtained when varying the bandwidth, except for a significant difference for Quechua ethnicity by female spouse eligibility, although only for bandwidth value 10. They can be found in Figure A.1 in the Appendix.[12]

Since in our research design we focus on pension eligibility of one or both spouses in a household, we quantify in Table A.1 in the Appendix the extent of endogenous sorting, which may alter couple formation, by regressing a dummy equal to 1 if the (female) male spouse is not observed in a household on the RDD specification for pension eligibility for the other spouse, after adding to our dataset on households with both spouses observations on single-spouse households. Non-significant coefficients of the eligibility dummies in Table A.1 suggest that endogenous sorting is not a major threat to our research design.[13]

In short, our findings confirm that that self-selection into the pension in households in which no spouse is eligible is not a concern empirically and, in addition, that those households in which at least one spouse is eligible and those in which no spouse is eligible are similar in a rich set of predetermined characteristics.

5 Results

We firstly show estimates of the pension take-up, of poverty and labour supply by spouses' eligibility by gender and also separately by our proxy for household wealth in section 5.1. Secondly, we show estimates of the pension eligibility effect on household composition in section 5.2. In addition, we discuss the robustness of our main results to varying the age bandwidth in section 5.3 and quantify ant(posti)icipation and placebo effects in section 5.4. Finally, in section 5.5 we compare and contrast our main results with additional ones obtained by using a different specification in which we estimate the effect of eligibility for the youngest and eldest spouse in a household, thus disregarding gender effects.

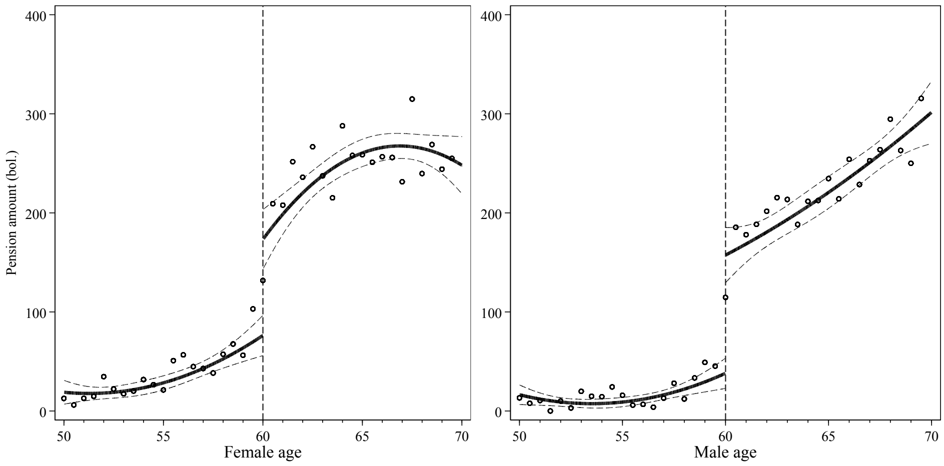
[12] Results obtained using polynomial plots of baseline characteristics, estimated separately by whether individuals are at least 60 and by gender are in line with those in Table 2 and can be found in Figure A.2 in the Appendix.

[13] We ran the same regressions as in Table A.1 except using as outcomes dummies for whether the marital status was widow, divorced or separated to assess whether any of them was induced by eligibility and found no evidence in support of this. We do not report these estimates although they are available upon request.

5.1 Take-up, poverty and consumption

We start by graphically assessing the effect of pension eligibility on take-up by gender. Figure 5 shows fitted RDD second order polynomials (thick continuous lines) and confidence intervals (dashed lines) of the total pension amount received by a household, obtained separately by whether each spouse was 60 or older. In addition, we show, as circles, mean values by age.

Figure 5: Pension amount received monthly in a household by spouse's gender



Two facts can be observed when looking at Figure 5. First, being individually eligible for the pension discontinuously increases the amount received by the household, as the non-overlapping confidence intervals at the cutoff show, separately for each spouse. Second, a small although non-zero amount is received in households on the left hand-side of the age 60 cutoff, with the amount being greater for females. This suggests that more than one spouse may be eligible for the pension in a household since non-compliance is very low, as shown in Table 1.[14]

Hence, failing to account for the possibility that more than one spouse in a household can be eligible for a pension can lead to biased RDD estimates, even though the research design is valid at the individual level. We thus formally quantify take-up by eligibility status separately by spouse's gender estimating equation (1) with the pension amount as dependent variable.

[14]The results are similar if we measure take-up using a dummy equal to one if a spouse received the pension and 0 otherwise, as shown in Table A.2 in the Appendix.

We report the estimates in Table 3. The first two columns show estimates of the pension amount, respectively excluding and adding as controls the baseline characteristics described in section 4.1. We report estimates by spouse gender (D_F and D_M), as well as the joint effect ($JE = D_F + D_M + D_{FM}$), together with the p-value of the null hypothesis of no joint effect.

In Table 3 we also report estimates of a model allowing for heterogeneous take-up by whether a household paid for health insurance in columns (3) and (4), respectively excluding and adding controls. In addition to estimated parameters associated to eligibility dummies by spouse (D_F and D_M), we report estimates of the health insurance dummy (HI), of its interactions with the eligibility dummies and of the following p-values: of the test that spouses' joint eligibility effect (JE) is zero and of the test that the heterogeneous joint effect for wealthy households, i.e. $HJE = D_F * HI + D_M * HI + D_{FM} * HI + HI$, is zero.

Table 3: Pension amount and poverty

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Pension amount				Poverty prob.			
D_F	138.039***	139.122***	180.312***	177.337***	0.017	0.063	-0.021	0.085
	(28.582)	(27.755)	(27.814)	(28.070)	(0.188)	(0.189)	(0.212)	(0.224)
D_M	108.516***	109.572***	108.318***	111.198***	-0.049	0.016	0.067	0.079
	(24.739)	(24.377)	(28.902)	(28.650)	(0.162)	(0.142)	(0.199)	(0.183)
D_{FM}	-48.065	-42.138	-110.180*	-100.328*	0.145	-0.006	0.069	-0.128
	(51.527)	(49.541)	(59.414)	(58.072)	(0.257)	(0.246)	(0.294)	(0.292)
1+ health ins. (HI)		-11.351	42.760*	32.857		-0.143***	-0.486**	-0.259
		(6.917)	(25.199)	(25.905)		(0.045)	(0.231)	(0.210)
$D_F * HI$			-147.626**	-136.859**			-0.127	-0.296
			(68.925)	(64.183)			(0.344)	(0.345)
$D_M * HI$			5.517	0.225			-0.332	-0.252
			(47.413)	(47.201)			(0.299)	(0.263)
$D_{FM} * HI$			298.891***	277.779***			0.473	0.729
			(95.884)	(92.343)			(0.455)	(0.463)
$JE = D_F + D_M + D_{FM}$	198.490***	206.557***	178.449***	188.207***	0.113	0.073	0.115	0.037
P-value (JE=0)	0.000	0.000	0.000	0.000	0.517	0.663	0.577	0.859
$HJE = HI(1 + D_F + D_M + D_{FM})$			199.542***	174.003***			-0.473**	-0.078
P-value (HJE=0)			0.000	0.001			0.040	0.769
Covariates	No	Yes	No	Yes	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048	1,048	1,048	1,048	1,048
Mean value (55-59)	6.426				0.460			
S. D. (55-59)	36.149				0.499			

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Columns (1) and (2) in Table 3 show that the Renta Dignidad pension amount is significantly higher when the female spouse is eligible for the pension (about 140 bol.), when the male is (about 110 bol.) and when both spouses are (about 200 bol.). Compared to the ideal situation of full compliance, we observe that households receive, on average, half the amount expected. This is consistent with a probability of receiving the pension varying between 49-66 percentage points, as shown in Table A.2 in the Appendix. When focusing on the heterogeneous take-up of the pension by health insurance in columns (3) and (4), we observe that take-up is significantly lower when only the female spouses is eligible in wealthy households, although this results is not robust when we vary the bandwidth, as shown in Figure A.3 in the Appendix; take-up is also significantly higher in households with a health insurance and two eligible spouses (about 280 bol.) and robust to changing the bandwidth.

In addition, since pension eligibility may lead to a mechanical increase in household income, Table 3 shows in the last four columns estimates of the pension eligibility effect on the probability that a household is poor, with poverty measured using a dummy equal to 1 if income is smaller than the poverty line. Focusing on columns (7) and (8), when both spouses are eligible the probability that income in a household is below the poverty line is lower in households with a health insurance. However, the effect loses significance when including in the regression baseline characteristics and it is not robust to varying the bandwidth, as shown in Table A.3 in the Appendix.

Table 4: Food consumption

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Total food expenditure				Share food exp. over total expenditures			
D_F	50.868	26.939	85.674	77.085	-0.084	-0.079	-0.087	-0.021
	(102.325)	(106.688)	(135.357)	(138.403)	(0.104)	(0.092)	(0.126)	(0.117)
D_M	-20.425	-26.187	-35.561	-33.725	-0.052	-0.031	-0.048	-0.048
	(83.680)	(88.760)	(111.683)	(119.698)	(0.093)	(0.082)	(0.111)	(0.108)
D_{FM}	-61.417	-53.628	-48.927	-87.101	0.154	0.066	0.155	-0.009
	(127.801)	(131.624)	(163.467)	(165.406)	(0.149)	(0.130)	(0.177)	(0.160)
1+ health ins. (HI)		-14.158	6.053	18.135		-0.023	-0.156	-0.019
		(18.791)	(117.156)	(123.669)		(0.024)	(0.119)	(0.110)
$D_F * HI$			-51.533	-108.504			0.011	-0.194
			(152.648)	(163.572)			(0.170)	(0.156)
$D_M * HI$			104.846	90.517			0.019	0.058
			(144.077)	(159.304)			(0.177)	(0.160)
$D_{FM} * HI$			-185.227	10.571			0.047	0.396
			(209.969)	(218.230)			(0.285)	(0.274)
$JE = D_F + D_M + D_{FM}$	-30.975	-52.876	1.186	-43.741	0.018	-0.043	0.020	-0.077
P-value $(_{JE=0})$	0.748	0.587	0.992	0.734	0.858	0.635	0.868	0.468
$HJE = HI(1 + D_F + D_M + D_{FM})$			-125.861	10.719			-0.079	0.241
P-value $(_{HJE=0})$			0.283	0.929			0.673	0.215
Covariates	No	Yes	No	Yes	No	Yes	No	Yes
N	1,045	1,045	1,045	1,045	1,045	1,045	1,045	1,045
Mean value (50-59)		296.343					0.441	
S. D. (50-59)		353.385					0.319	

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Finally, we quantify the extent to which pension eligibility induced higher food consumption, which is typically elastic to changes in income, particularly in a developing country setting. Table 4 reports estimates obtained using as outcome food consumption, measured both in bolivianos and as percentage of total expenditure in a household. The eligibility effect on food consumption is not significant independently of whether the outcome is measured in bolivianos or in percentage of total expenditure, as shown in columns (1)-(4) and (5)-(8) respectively. When we look at the point estimates of regressions using consumption in bolivianos as outcome, we find that the eligibility effect for female spouses is positive while it is negative for male spouses; when the outcome is consumption as percentage of total expenditure, estimates are negative both for the female and male spouse.[15]

5.2 Household composition

In this section we report evidence of the pension eligibility effect by focusing on co-residence of elderly spouses' grandchildren as one or more pensions may give grandparents additional resources and decrease the opportunity cost of childcare. Results on co-residence of elderly spouses' adult children or of extended family members are reported in the Appendix as they are not significant. Table 5 shows estimates of the pension eligibility effect on the probability that spouses' grandchildren live with them, as well as on educational expenditure. Columns (1) and(2) show that the probability that grandchildren co-reside with their grandparents is higher when both spouses are eligible, being significant (about 50 pp or 280%). While this result becomes weakly significant when we allow for an heterogeneous effect by wealth in columns (3) and (4), the joint effect of eligibility for both spouses reported at the bottom of the table is significant for all specifications.

Since grandchildren are typically in schooling age, we also assess whether

[15]The small difference in the number of observations reported in Table 4 relative to the other tables reporting results is due to trimming three very suspicious outliers for food consumption whose values is about 100,000 bolivianos. Results on the pension eligibility effect on extreme poverty and labour supply are not reported as they are either not significant or somewhat significant but not robust to varying the bandwidth. However, they can be found in Table A.3-A.5 in the Appendix. We, instead, do not report results obtained using as outcomes consumption of durables or total consumption as they are not significant although they are available upon request.

pension eligibility induces changes in educational expenditure. Columns (5) to (8) in Table 5 show no significant effect on educational expenditure and no significant difference by household wealth. Note that the number of observations in columns (5) to (8) is lower due to the higher frequency of missing values for the variable measuring educational expenditure. Overall, the evidence in Table 5 suggests that eligible grandparents' higher investment is in time spent with their grandchildren and in bearing the associated costs, rather than in educational expenditure. This evidence is complemented by the one showing no change in the probability of co-residence for elderly spouses' adult children, which can be found in Table A.6 and A.7 in the Appendix. The reason is that, contrary to related studies on South Africa showing that young adults migrate when their parents become eligible for a pension, we document that the intergenerational effect induced by the Bolivian pension on non-beneficiaries affects only eligibles' grandchildren.

Table 5: Co-residing grandchildren

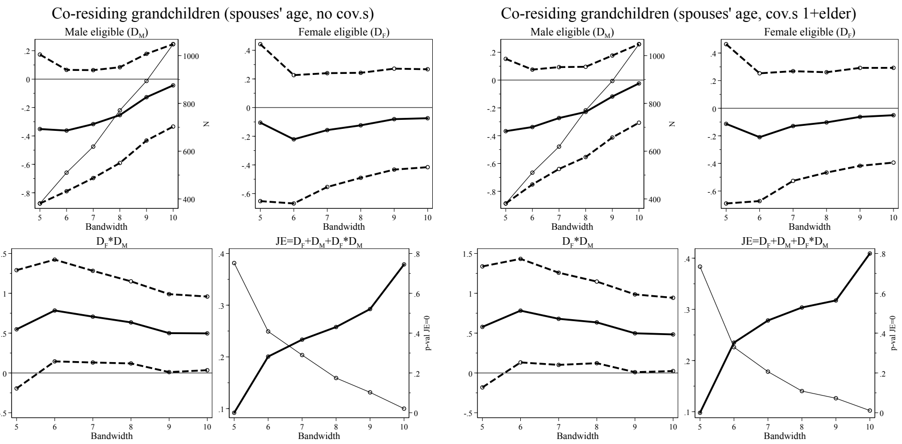
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Co-residing grandchildren				Educational expenditure			
D_F	-0.074	-0.051	0.019	0.030	52.466	46.921	84.231	63.312
	(0.174)	(0.175)	(0.195)	(0.198)	(74.233)	(67.291)	(83.513)	(76.280)
D_M	-0.044	-0.024	0.002	0.038	2.977	-9.748	-11.187	-16.537
	(0.148)	(0.144)	(0.170)	(0.167)	(39.747)	(38.706)	(44.996)	(43.469)
D_{FM}	0.497**	0.484**	0.499*	0.487*	-12.085	17.598	-39.266	15.417
	(0.236)	(0.234)	(0.262)	(0.264)	(84.438)	(75.480)	(91.534)	(82.960)
1+ health ins. (HI)		0.023	0.263	0.299		9.131	91.604	48.039
		(0.040)	(0.232)	(0.225)		(16.416)	(71.982)	(70.915)
$D_F * HI$			-0.245	-0.191			-176.311	-97.638
			(0.403)	(0.401)			(111.920)	(106.478)
$D_M * HI$			-0.229	-0.274			29.414	21.514
			(0.320)	(0.309)			(96.063)	(95.254)
$D_{FM} * HI$			-0.138	-0.173			185.160	50.918
			(0.542)	(0.536)			(165.542)	(157.315)
$JE = D_F + D_M + D_{FM}$	0.379**	0.409**	0.520***	0.555***	43.357	54.772	33.778	62.193
P-value ($JE=0$)	0.021	0.011	0.004	0.002	0.291	0.161	0.415	0.130
$HJE = HI(1 + D_F + D_M + D_{FM})$			-0.348	-0.339			129.867	22.833
P-value ($HJE=0$)			0.226	0.235			0.212	0.816
Covariates	No	Yes	No	Yes	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048	990	990	990	990
Mean value (55-59)	0.171				208.361			
S. D. (55-59)	0.377				438.602			

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

5.3 Robustness to varying age bandwidth

We now assess the robustness of our main result, i.e. a higher probability of observing co-residing grandchildren in households with both spouses eligible for a pension, to considering only households whose spouses' age are

Figure 6: Robustness of eligibility effect to varying the bandwidth



as close as possible to the 60 cutoff. The smallest feasible bandwidth is 5 since with smaller values our model, which has a high number of interaction terms, tends to be misspecified due to the low number of observations.

Figure 6 shows how estimates of the pension eligibility effect by spouse's gender on the probability of co-residing grandchildren, measured on the vertical axis, vary with the bandwidth for male and female spouse age, measured on the horizontal axis, using the same econometric specification as in equation (1). We report estimates of the eligibility effect excluding and adding controls, respectively in the left-hand side and right-hand side panel in Figure 6. Overall, our main results are unchanged when we vary the age bandwidth, except for bandwidth value 5 when our estimates are weakly significant, which is at least in part due to the fact that the number of observations is smaller than 400, i.e. less than 40% of the number of observations when the bandwidth value is 10. Estimates for poverty, consumption, labor supply and for the probability of observing co-residing adult children, along with estimates of the heterogeneous effect by wealth are reported in Figure A.3 in the Appendix as they are not significant.[16]

In addition, we assess whether our main results are unchanged once we account for the potential confounding effect of Bonosol, the non-contributory

[16]Our main results are unchanged when dropping observations of households receiving a contributory pension, as the eligibility cutoff age for females coincides with the one for Renta Dignidad, and when dropping households in which non-core members received a Renta Dignidad pension. These results are not reported although they are available upon request.

pension that was in place until 2007 with an age eligibility cutoff set at 65. The reason is that when we consider bandwidth 5 our data sample contains, for example, individuals who were 64 in 2007, did not receive Bonosol then and started receiving Renta Dignidad in 2008. By contrast, it does not contain individuals who were 65 or older and therefore they first received Bonosol until it was discontinued and then started receiving Renta Dignidad in 2008. To assess whether our main results vary when we drop from the dataset only individuals who were 65 or older in 2008 or 2009, as they may have been eligible for Bonosol, we set the left-hand side bandwidth to 10, i.e. consider households in which both spouses are in the 50-60 age range, and decrease the right-hand side bandwidth from 10 to 5. Estimates, which are in line with our main results, are reported in Figure A.4 in the Appendix.

5.4 Ant(post)icipation and placebo effects

In this section we, first, quantify *anticipation* effects. An example is a decrease in poverty in a household in which one spouse is close to turn 60 and decides to borrow in advance of becoming eligible for the Renta Dignidad pension. We also quantify *posticipation*, i.e. lagged, effects to assess if household decisions respond with a lag to the change in eligibility status. Second, we quantify *placebo* effects, i.e. we test if households modify, for example, consumption or labour supply when one or more spouses turn 60 even before the Renta Dignidad pension was enacted. This is particularly important to quantify whether our main results are clean estimates of the Renta Dignidad pension for females or whether they are confounded by the contributory pension effect since the contributory pension age cutoff for females is 60.

We start by quantifying anticipation and posticipation effects. We do so by considering the same RDD empirical specification we used to obtain our main results in equation (1), except that we now use as cutoff a small(great)er age value than 60 for ant(post)icipation effects, while age is rescaled accordingly. Figure 7 reports estimates of several regressions for different age cutoff values in a compact form as follows: estimates of the female, male spouse and spouses' joint pension eligibility effect on the probability of observing co-residing grandchildren in a household are measured on the vertical axis for different age cutoff values measured on the horizontal axis.

Figure 7: Ant(post)icipation effects of pension eligibility

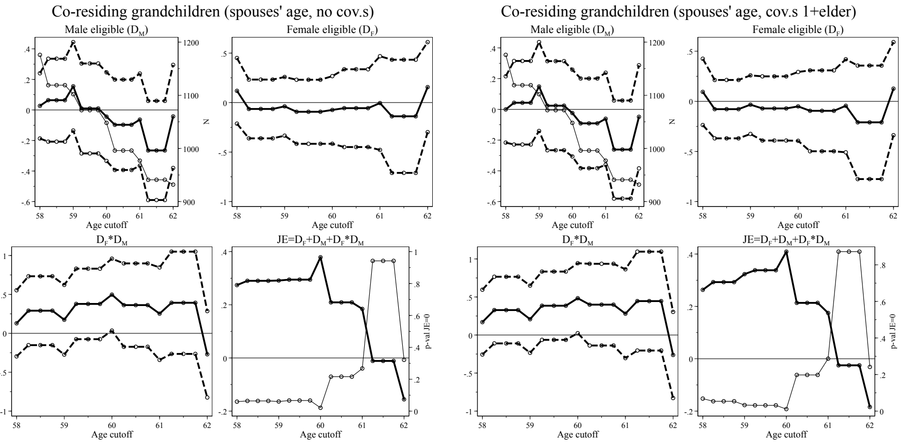


Figure 7 shows for an age cutoff value of 60 the effect of eligibility on the probability of observing co-residing grandchildren, which is positive and significant only when both spouses are eligible and was reported in Table 5. For values of the age cutoff smaller than 60, estimates quantify the anticipation effect and vice versa a posticipation effect for values greater than 60. Figure 7 shows that the only significant effect is at 60, thus offering evidence in support of the absence of anticipation and of lagged effects.[17]

We now turn to quantifying placebo effects. We do this by considering, as we did for ant(post)icipation effects, the same RDD empirical specification we used to obtain our main results in equation (1), except that we now estimate it using data from the 2006 and 2007 waves of the Bolivian household survey. By doing so we can test whether our main result, i.e. a higher probability of observing co-residing grandchildren in households in which both spouses are eligible, is truly due to the Renta Dignidad pension eligibility, in which case we should find that estimates of the effect that both spouses turn 60 in 2006-2007 are not significant, rather than to the potentially confounding demographic effect of turning 60 or to the contributory pension threshold for females which is set at 60.

[17]Estimates of the ant(post)icipation effect on the following outcomes: poverty dummy, food consumption, labour supply by at least one spouse dummy and dummies for co-residing adult children and for grandchildren can be found in Figure A.5 in the Appendix.

Figure 8: Placebo effects of pension eligibility using 2006-2007 data

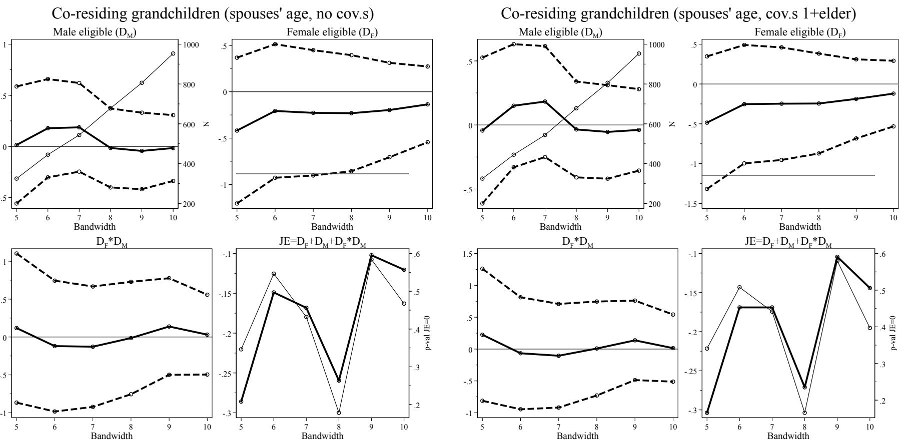


Figure 8 reports on the vertical axis estimates obtained with as outcome the dummy equal to 1 if co-residing grandchildren are observed in a household, as in Figure 7. The figure reports estimates from regressions with different bandwidth values on the horizontal axis. We first focus on bandwidth value 10, which is the one we used to obtain our main results. Figure 8 shows that the effect of turning 60 is not significant neither for each spouse separately nor for both spouses jointly, which suggests that our main results are truly due to the pension eligibility rather than to spouses turning 60. When we look at the estimates obtained using different bandwidth values the results are unchanged.[18]

5.5 Eligibility effect disregarding gender

In addition to our main specification which studies the effect of pension eligibility by gender, we also study the effect of eligibility of at least one spouse in a household, to critically assess what results we would forego by not accounting for spouses' eligibility by gender. This is done by following as guidance the specification used in Escobar Loza *et al.* (2013). They define as treated those households in which the eldest individual was eligible for the pension, which implies that at least an individual is eligible as younger individuals in the households may also be eligible; controls are

[18]Estimates of the placebo effect on the following outcomes: poverty dummy, consumption, labour supply by at least one spouse dummy and dummies for co-residing adult children and for grandchildren can be found in Figure A.6 in the Appendix.

defined as those households with no individual eligible since the eldest is not eligible, which implies that younger individuals are not eligible either. The effect is estimated by using a RDD in which the running variable is the maximum age. The difference in our specification not accounting for gender differences relative to the one in Escobar Loza *et al.* (2013) is that we focus on spouses rather than on all individuals in a household since in 99% of households the eldest individual coincides with the eldest spouse when the age range is within 50-70.

Table 6: Effect of eligibility of the eldest spouse in a household on grandchildren's co-residence and educ. expenditure

	Co-residing grandchildren		Educ. expenditure (bol.)	
DMaxA	0.060	0.064	-124.566	20.334
	(0.050)	(0.068)	(240.008)	(57.374)
(MaxA-60)	0.010**	-0.000	-102.494	-5.658
	(0.005)	(0.018)	(106.469)	(12.176)
DMaxA*(MaxA-60)	-0.006	0.009	155.655	-7.809
	(0.009)	(0.025)	(172.306)	(18.631)
N	1,551	831	1,551	831
Mean value (incl. HH)	0.167	0.171	1264.487	1174.655
S.d. (incl. HH)	0.373	0.377	31493.721	30331.384
Covariates	Yes	Yes	Yes	Yes
Bandwidth	10	5	10	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 6 reports estimates of the effect of eligibility for the eldest spouse in a household, using a RDD with the maximum age between spouses as running variable. Table 6 shows that the effect on the probability of observing co-residing grandchildren is positive but not significant. The effect on educational expenditure is not significant either and the sign varies with the bandwidth used. We do not report estimates obtained using the eldest individual in a household as they are little different although they are available upon request.[19]

[19]Table A.8 in the Appendix reports estimates of the eligibility effect for the eldest spouse in a household on take-up, food expenditure and labour supply, with no estimate being significant. Table A.8 also reports estimates of the effect that both spouses are eligible, which we obtained using a RDD with the minimum age between spouses as running variable. The only significant estimate is a negative effect on labour supply by female spouses.

6 Discussion

Our RDD estimates show that, although eligibility for the Renta Dignidad pension separately by spouse gender increases take-up, poverty incidence is not lower in households with at least one spouse eligible for the pension. In addition, they show that the main behavioural response of eligibility is a change in household composition, with two eligible grandparents being more likely to live with one or more grandchildren. Our paper complements the growing number of related studies on social pensions in Mexico, in Latin American countries and in South Africa by shedding light on the role of eligibility by spouse gender, which has been little studied to the best of our knowledge and, therefore, has not been linked to the impact on (non-)beneficiaries.

Differently from two closely related studies on the same pension in Bolivia, ours does not find that in the short-run the pension decreases poverty nor female labour supply. Relative to Escobar Loza *et al.* (2013), when we replicate their setup which focuses on eligibility by gender of the eldest household member, we find that spouses, separately by gender, are in 99% of the households the eldest members, that the impact on poverty is non-significant, in line with our results accounting by spouses' gender, and that the impact on co-residing grandchildren is also not significant, differently from the positive effect we find when both spouses are eligible. This is partly due to accounting for gender differences although it may be also due to the fact that our datasets are not the same since Escobar Loza *et al.* (2013) obtain a dataset of households with the eldest individual in the 55-65 age range and with a greater number of observations than in the household survey data we used. Relative to Hernani-Limarino and Mena (2015), the main differences are, first, that they focus on eligibility of the eldest individual in a household, like Escobar Loza *et al.* (2013) and, second, that they also use data on the previous pension, Bonosol, to implement a refinement of a difference in difference design.

From a policy viewpoint, our results suggest that in a country with widespread poverty, such as Bolivia, a social pension seems to lead to potential intergenerational "spillovers" onto beneficiaries' extended family which may have not been anticipated at the policy design stage. This suggests that a thorough analysis of the impact of a social pension encompasses both gender roles and indirect behavioural responses to the pension income shock, rather than just poverty rates or labour supply as main proxies for well-being.

A few caveats in our analysis should be noted. First, behavioural responses to the pension are observed shortly after individuals have become eligible, i.e. in the short-run, while less is known about their long-run impact. This is due to the cross-sectional nature of the survey data which does not allow us to follow households over time. Second, the eligibility effect has been estimated around the age 60 cutoff while it may be different for eligible individuals at different ages. Third, in our analysis we estimated the intent-to-treat effect as take-up is endogenous, which makes us mainly interested in the sign of our estimates rather than their magnitude.

Our paper paves the way for future related research. First, it would be valuable to quantify whether benefits accruing to households with eligible spouses also indirectly "spillover" onto households in their social network or located nearby but with no eligible spouse since the focus of social pension studies is almost entirely on eligibles. Second, it would be valuable to test whether and to what extent the timing of information disclosure about social pensions enactment play a relevant role in explaining take-up and behavioural decisions of households with (no) eligible members. Third, in addition to married couples, extending our focus to widows may be helpful to learn about the eligibility effect in single-spouse households and also about the loss-of-eligibility effect in the event of the death of an eligible spouse in a household. Assessing the empirical relevance of these aspects will help policy-makers anticipating their potential impact on the subsequent analysis of a social pension effect.

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Online Appendix

Table A.1: Missing spouse by pension eligibility for the other spouse

	Missing female spouse dummy				Missing male spouse dummy			
DM	-0.019	-0.018	-0.016	-0.011				
	(0.044)	(0.044)	(0.060)	(0.060)				
(AM-60)	0.009**	0.012***	-0.006	-0.008				
	(0.005)	(0.004)	(0.016)	(0.015)				
DM*(AM-60)	0.002	0.001	0.021	0.027				
	(0.008)	(0.008)	(0.021)	(0.021)				
DF					-0.023	-0.010	0.007	-0.005
					(0.051)	(0.050)	(0.071)	(0.069)
(AF-60)					0.004	0.010*	0.004	0.010
					(0.006)	(0.006)	(0.016)	(0.016)
DF*(AF-60)					0.014	0.010	0.007	0.012
					(0.009)	(0.009)	(0.025)	(0.025)
Constant	0.215***	0.276***	0.189***	0.240***	0.338***	0.284***	0.319***	0.278***
	(0.030)	(0.042)	(0.045)	(0.060)	(0.034)	(0.046)	(0.048)	(0.066)
N	1,874	1,874	970	970	1,914	1,914	926	926
Mean value (incl. HH)	0.176	0.176	0.205	0.205	0.304	0.304	0.313	0.313
S.d. (incl. HH)	0.381	0.381	0.404	0.404	0.460	0.460	0.464	0.464
Covariates	0	1	0	1	0	1	0	1
Bandwidth	10	10	5	5	10	10	5	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.2: Pension take-up probability by spouse gender

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Females				Males			
D_F	0.660***	0.655***	0.866***	0.858***	0.064	0.058	0.073	0.058
	(0.121)	(0.118)	(0.101)	(0.099)	(0.067)	(0.075)	(0.087)	(0.090)
D_M	-0.016	-0.024	0.051	0.059	0.607***	0.594***	0.498***	0.503***
	(0.068)	(0.069)	(0.055)	(0.058)	(0.106)	(0.102)	(0.128)	(0.125)
D_{FM}	-0.090	-0.058	-0.389**	-0.361**	-0.173	-0.146	-0.211	-0.175
	(0.167)	(0.163)	(0.167)	(0.164)	(0.164)	(0.160)	(0.199)	(0.195)
1+ health ins. (HI)		-0.007	0.236*	0.191		-0.045**	-0.023	-0.046
		(0.025)	(0.123)	(0.120)		(0.022)	(0.028)	(0.035)
$D_F * HI$			-0.690**	-0.690**			-0.066	-0.010
			(0.334)	(0.320)			(0.087)	(0.100)
$D_M * HI$			-0.203	-0.221			0.385**	0.375**
			(0.169)	(0.170)			(0.174)	(0.169)
$D_{FM} * HI$			1.153***	1.132***			0.464*	0.372
			(0.386)	(0.371)			(0.240)	(0.240)
$JE = D_F + D_M + D_{FM}$	0.554***	0.573***	0.529***	0.556***	0.498***	0.506***	0.361***	0.385***
P-value ($_{JE=0}$)	0.000	0.000	0.000	0.000	0.000	0.000	0.006	0.003
$HJE = HI(1 + D_F + D_M + D_{FM})$			0.496***	0.413***			0.760***	0.691***
P-value ($_{HJE=0}$)			0.001	0.006			0.000	0.000
Covariates	No	Yes	No	Yes	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048	1,048	1,048	1,048	1,048
Mean value (50-59)	0.026				0.008			
S. D. (50-59)	0.160				0.089			

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.3: Extreme poverty prob.

	(1)	(2)	(3)	(4)
D_F	-0.333**	-0.316**	-0.312*	-0.218
	(0.144)	(0.135)	(0.179)	(0.173)
D_M	-0.115	-0.058	-0.019	0.020
	(0.153)	(0.133)	(0.198)	(0.177)
D_{FM}	0.492**	0.373*	0.437*	0.225
	(0.213)	(0.194)	(0.263)	(0.243)
1+ health ins. (HI)		-0.114***	-0.201	-0.012
		(0.032)	(0.179)	(0.174)
$D_F * HI$			-0.078	-0.303
			(0.244)	(0.248)
$D_M * HI$			-0.260	-0.246
			(0.265)	(0.234)
$D_{FM} * HI$			0.071	0.472
			(0.341)	(0.334)
$JE = D_F + D_M + D_{FM}$	0.044	-0.002	0.106	0.027
P-value ($_{JE=0}$)	0.774	0.989	0.579	0.887
$HJE = HI(1 + D_F + D_M + D_{FM})$			-0.468***	-0.090
P-value ($_{HJE=0}$)			0.001	0.571
Covariates	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048
Mean value (50-59)		0.299		
S. D. (50-59)		0.458		

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.4: Labor supply (1+ spouse)

	(1)	(2)	(3)	(4)
D_F	0.017	0.045	-0.068	-0.038
	(0.084)	(0.086)	(0.043)	(0.047)
D_M	-0.054	-0.005	-0.073*	-0.058
	(0.066)	(0.059)	(0.040)	(0.044)
D_{FM}	-0.140	-0.180	0.045	0.025
	(0.129)	(0.120)	(0.073)	(0.078)
1+ health ins. (HI)		-0.003	-0.116	-0.084
		(0.023)	(0.108)	(0.104)
$D_F * HI$			0.168	0.123
			(0.270)	(0.272)
$D_M * HI$			0.000	0.004
			(0.178)	(0.176)
$D_{FM} * HI$			-0.630	-0.586
			(0.432)	(0.434)
$JE = D_F + D_M + D_{FM}$	-0.177**	-0.141*	-0.096	-0.072
P-value ($_{JE=0}$)	0.046	0.076	0.124	0.294
$HJE = HI(1 + D_F + D_M + D_{FM})$			-0.578*	-0.543*
P-value ($_{HJE=0}$)			0.060	0.076
Covariates	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048
Mean value (50-59)		0.950		
S. D. (50-59)		0.219		

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.5: Labor supply

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Females				Males			
D_F	-0.210	-0.194	-0.371*	-0.324*	0.026	0.061	-0.048	0.001
	(0.178)	(0.169)	(0.192)	(0.185)	(0.113)	(0.118)	(0.119)	(0.124)
D_M	-0.066	-0.022	-0.294*	-0.267	-0.024	0.032	-0.027	-0.012
	(0.150)	(0.144)	(0.174)	(0.170)	(0.105)	(0.104)	(0.125)	(0.130)
D_{FM}	0.197	0.147	0.483*	0.391	-0.230	-0.291*	-0.061	-0.119
	(0.242)	(0.234)	(0.274)	(0.270)	(0.167)	(0.161)	(0.167)	(0.170)
1+ health ins. (HI)		-0.071	-0.606***	-0.557***		0.009	-0.067	-0.036
		(0.051)	(0.193)	(0.192)		(0.030)	(0.159)	(0.155)
$D_F * HI$			0.414	0.264			0.141	0.025
			(0.406)	(0.396)			(0.294)	(0.294)
$D_M * HI$			0.695**	0.691**			-0.039	-0.041
			(0.293)	(0.291)			(0.232)	(0.227)
$D_{FM} * HI$			-0.872	-0.607			-0.622	-0.480
			(0.557)	(0.554)			(0.470)	(0.470)
$JE = D_F + D_M + D_{FM}$	-0.079	-0.069	-0.181	-0.201	-0.228	-0.198	-0.135	-0.130
P-value (JE=0)	0.621	0.659	0.286	0.244	0.086	0.115	0.378	0.395
$HJE = HI(1 + D_F + D_M + D_{FM})$			-0.369	-0.209			-0.588*	-0.532
P-value (HJE=0)			0.237	0.515			0.071	0.101
Covariates	No	Yes	No	Yes	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048	1,048	1,048	1,048	1,048
Mean value (50-59)	0.624				0.922			
S. D. (50-59)	0.485				0.269			

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.6: Co-residing adult children of elderly couples

	(1)	(2)	(3)	(4)
D_F	-0.225	-0.231	-0.239	-0.281
	(0.172)	(0.159)	(0.192)	(0.190)
D_M	-0.113	-0.126	-0.087	-0.098
	(0.140)	(0.143)	(0.173)	(0.179)
D_{FM}	0.386	0.423*	0.312	0.396
	(0.241)	(0.230)	(0.276)	(0.273)
1+ health ins. (HI)		-0.024	-0.100	-0.152
		(0.042)	(0.192)	(0.196)
$D_F * HI$			-0.046	0.074
			(0.378)	(0.337)
$D_M * HI$			-0.158	-0.139
			(0.291)	(0.288)
$D_{FM} * HI$			0.352	0.132
			(0.550)	(0.528)
$JE = D_F + D_M + D_{FM}$	0.048	0.066	-0.014	0.018
P-value (JE=0)	0.760	0.680	0.938	0.922
$HJE = HI(1 + D_F + D_M + D_{FM})$			0.048	-0.085
P-value (HJE=0)			0.885	0.808
Covariates	No	Yes	No	Yes
N	1,048	1,048	1,048	1,048
Mean value (50-59)	0.795			
S. D. (50-59)	0.404			

Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.7: Share of educational expenditure over total expenditure

	(1)	(2)	(3)	(4)
D_F	0.007	0.005	-0.067	-0.088
	(0.105)	(0.098)	(0.117)	(0.110)
D_M	-0.099	-0.101	-0.104	-0.109
	(0.066)	(0.064)	(0.083)	(0.080)
D_{FM}	0.041	0.074	0.104	0.171
	(0.118)	(0.110)	(0.131)	(0.123)
1+ health ins. (HI)		0.018	0.064	0.013
		(0.025)	(0.110)	(0.109)
$D_F * HI$			0.369	0.427*
			(0.279)	(0.251)
$D_M * HI$			-0.007	-0.004
			(0.133)	(0.130)
$D_{FM} * HI$			-0.295	-0.426
			(0.310)	(0.284)
$JE = D_F + D_M + D_{FM}$	-0.051	-0.021	-0.066	-0.026
P-value ($_{JE=0}$)	0.439	0.745	0.405	0.743
$HJE = HI(1 + D_F + D_M + D_{FM})$			0.131	0.010
P-value ($_{HJE=0}$)			0.245	0.930
Covariates	No	Yes	No	Yes
N	991	991	991	991
Mean value (50-59)	0.282			
S. D. (50-59)	0.595			

Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.8: Effect of pension eligibility for 1+ spouse versus no (2 versus 0 or 1) using max(min) age

	Female take-up				Male take-up				Take-up amount (bolivianos)			
DMaxA	0.182***	0.208***			0.463***	0.457***			125.777***	129.567***		
	(0.040)	(0.052)			(0.039)	(0.052)			(10.707)	(13.491)		
(MaxA-60)	0.001	0.005			-0.001	-0.006			0.178	0.148		
	(0.002)	(0.007)			(0.001)	(0.004)			(0.506)	(1.547)		
DMaxA*(MaxA-60)	0.053***	0.031*			0.053***	0.065***			20.160***	17.770***		
	(0.008)	(0.019)			(0.006)	(0.018)			(1.966)	(5.018)		
DMinA			0.611***	0.507***			0.102*	-0.021			137.203***	93.788***
			(0.046)	(0.065)			(0.060)	(0.087)			(15.594)	(21.564)
(MinA-60)			0.006**	0.023**			0.033***	0.057***			7.747***	15.443***
			(0.003)	(0.010)			(0.006)	(0.019)			(1.391)	(4.163)
DMinA*(MinA-60)			0.029***	0.043**			0.010	0.010			7.358***	10.431
			(0.008)	(0.019)			(0.010)	(0.028)			(2.640)	(6.774)
Constant	0.008	0.023	-0.002	0.021	0.004	0.011	0.417***	0.539***	4.254	8.717	82.722***	110.037***
	(0.025)	(0.033)	(0.027)	(0.043)	(0.025)	(0.036)	(0.055)	(0.081)	(7.366)	(10.129)	(12.705)	(19.440)
N	1,551	831	1,276	628	1,551	831	1,276	628	1,551	831	1,276	628
Mean value (inel. HH)	0.300	0.336	0.332	0.377	0.290	0.330	0.308	0.334	93.063	102.763	99.071	107.085
S.d. (inel. HH)	0.458	0.472	0.471	0.485	0.454	0.470	0.462	0.472	136.827	140.559	140.078	143.908
Covariates	1	1	1	1	1	1	1	1	1	1	1	1
Bandwidth	10	5	10	5	10	5	10	5	10	5	10	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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	Poverty				Extreme poverty			
DMaxA	0.052	0.004			0.010	-0.001		
	(0.051)	(0.070)			(0.046)	(0.064)		
(MaxA-60)	-0.010*	0.003			-0.004	-0.011		
	(0.006)	(0.017)			(0.006)	(0.017)		
DMaxA*(MaxA-60)	0.004	0.006			-0.008	0.014		
	(0.010)	(0.024)			(0.008)	(0.023)		
DMinA			0.049	0.056			-0.070	-0.048
			(0.058)	(0.079)			(0.053)	(0.075)
(MinA-60)			-0.008	-0.012			0.001	-0.006
			(0.006)	(0.018)			(0.006)	(0.018)
DMinA*(MinA-60)			-0.001	0.007			-0.010	0.004
			(0.010)	(0.028)			(0.009)	(0.025)
Constant	0.734***	0.713***	0.721***	0.714***	0.505***	0.506***	0.511***	0.489***
	(0.050)	(0.069)	(0.054)	(0.076)	(0.046)	(0.067)	(0.052)	(0.078)
N	1,551	831	1,276	628	1,551	831	1,276	628
Mean value (incl. HH)	0.440	0.436	0.436	0.431	0.230	0.224	0.225	0.217
S.d. (incl. HH)	0.496	0.496	0.496	0.495	0.421	0.417	0.418	0.412
Covariates	1	1	1	1	1	1	1	1
Bandwidth	10	5	10	5	10	5	10	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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	Female spouse labour supply				Male spouse labour supply				1+ spouse labour supply			
DMaxA	-0.057	-0.122			-0.008	-0.045			-0.017	-0.053*		
	(0.056)	(0.077)			(0.034)	(0.046)			(0.024)	(0.029)		
(MaxA-60)	0.000	0.009			0.000	0.018			0.003	0.017**		
	(0.006)	(0.019)			(0.003)	(0.014)			(0.002)	(0.008)		
DMaxA*(MaxA-60)	-0.010	-0.000			-0.012*	-0.033*			-0.015**	-0.025**		
	(0.010)	(0.027)			(0.007)	(0.018)			(0.006)	(0.012)		
DMinA			-0.237***	-0.272***			-0.031	-0.023			-0.057	-0.065
			(0.063)	(0.084)			(0.048)	(0.063)			(0.037)	(0.048)
(MinA-60)			0.010	0.018			0.001	-0.000			0.003	-0.000
			(0.007)	(0.019)			(0.004)	(0.016)			(0.003)	(0.011)
DMinA*(MinA-60)			-0.007	-0.002			-0.011	-0.023			-0.011	-0.006
			(0.012)	(0.030)			(0.009)	(0.023)			(0.008)	(0.019)
Constant	0.658***	0.638***	0.739***	0.722***	0.947***	0.978***	0.956***	0.935***	0.985***	1.011***	0.973***	0.958***
	(0.053)	(0.073)	(0.057)	(0.079)	(0.029)	(0.039)	(0.039)	(0.054)	(0.019)	(0.022)	(0.028)	(0.037)
N	1,551	831	1,276	628	1,551	831	1,276	628	1,551	831	1,276	628
Mean value (incl. HH)	0.606	0.604	0.604	0.602	0.739	0.709	0.715	0.689	0.862	0.843	0.846	0.829
S.d. (incl. HH)	0.489	0.489	0.489	0.489	0.439	0.454	0.452	0.463	0.345	0.363	0.361	0.377
Covariates	1	1	1	1	1	1	1	1	1	1	1	1
Bandwidth	10	5	10	5	10	5	10	5	10	5	10	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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	Food expenditure (bolivianos)				Food expenditure (% total exp.)			
DMaxA	-4.434	0.550			0.030	0.040		
	(27.670)	(35.909)			(0.028)	(0.038)		
(MaxA-60)	1.519	4.331			0.005	0.002		
	(3.071)	(8.589)			(0.003)	(0.010)		
DMaxA*(MaxA-60)	7.923	1.857			0.000	-0.001		
	(5.769)	(13.888)			(0.005)	(0.014)		
DMinA			7.685	-0.179			-0.048	-0.028
			(34.578)	(44.074)			(0.037)	(0.048)
(MinA-60)			-0.385	4.930			0.012***	0.007
			(4.210)	(11.242)			(0.004)	(0.011)
DMinA*(MinA-60)			5.514	-5.990			0.001	-0.001
			(7.083)	(15.885)			(0.007)	(0.017)
Constant	272.542***	295.724***	267.734***	321.482***	0.579***	0.601***	0.645***	0.633***
	(26.158)	(35.603)	(32.091)	(43.496)	(0.027)	(0.037)	(0.033)	(0.048)
N	1,548	829	1,273	626	1,547	828	1,272	625
Mean value (incl. HH)	354.286	363.935	364.937	372.841	0.470	0.486	0.486	0.499
S.d. (incl. HH)	393.328	404.765	404.235	412.286	0.321	0.322	0.322	0.322
Covariates	1	1	1	1	1	1	1	1
Bandwidth	10	5	10	5	10	5	10	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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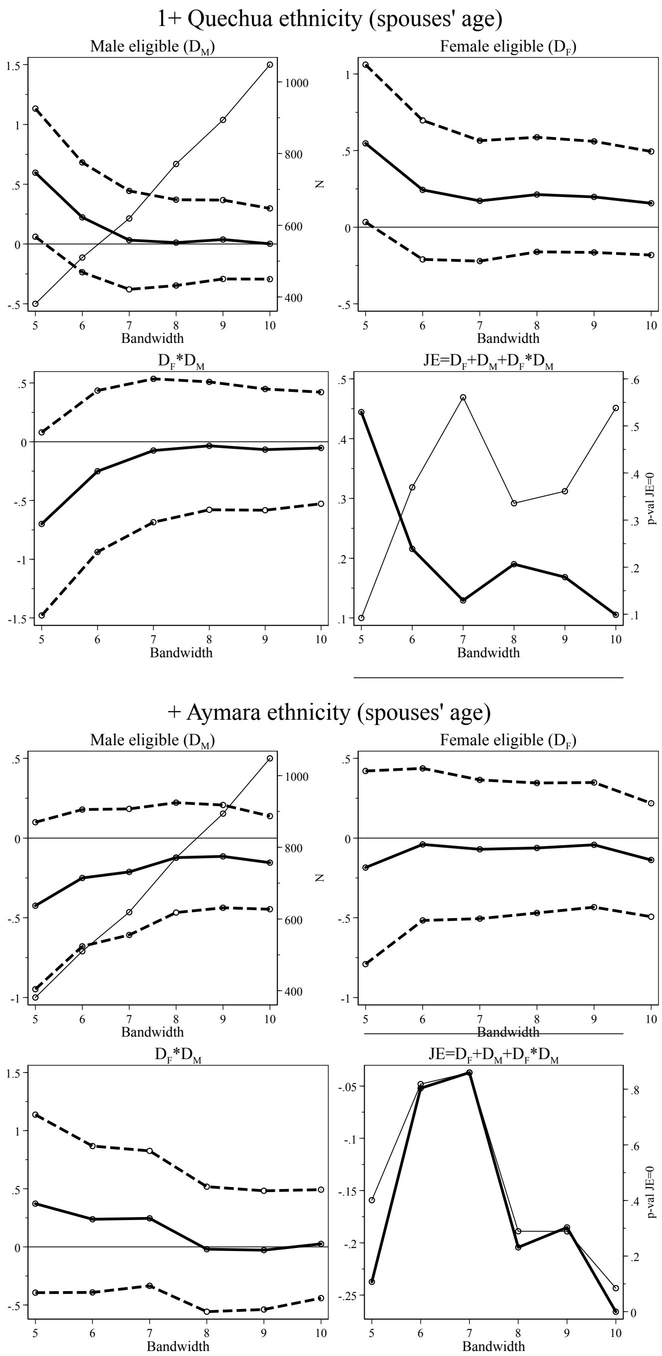
	Co-residing grandchildren				Educational expenditure (bolivianos)				Educational expenditure (% total exp.)			
DMaxA	0.060	0.064			-124.566	20.334			-0.002	0.000		
	(0.050)	(0.068)			(240.008)	(57.374)			(0.021)	(0.029)		
(MaxA-60)	0.010**	-0.000			-102.494	-5.658			-0.003	-0.002		
	(0.005)	(0.018)			(106.469)	(12.176)			(0.003)	(0.007)		
DMaxA*(MaxA-60)	-0.006	0.009			155.655	-7.809			-0.003	-0.005		
	(0.009)	(0.025)			(172.306)	(18.631)			(0.004)	(0.009)		
DMinA			0.091	0.073			416.935	84.134			-0.012	-0.017
			(0.062)	(0.085)			(434.370)	(76.267)			(0.022)	(0.029)
(MinA-60)			0.016***	0.028			-138.137	-17.479			-0.005*	-0.007
			(0.006)	(0.019)			(136.287)	(14.829)			(0.003)	(0.007)
DMinA*(MinA-60)			-0.035***	-0.059**			152.711	25.094			0.002	0.012
			(0.011)	(0.028)			(171.756)	(25.734)			(0.004)	(0.011)
Constant	0.231***	0.222***	0.284***	0.297***	-445.329	85.201	-610.336	-8.282	0.071***	0.072**	0.051**	0.035
	(0.044)	(0.060)	(0.054)	(0.074)	(551.898)	(55.709)	(701.068)	(47.702)	(0.020)	(0.029)	(0.020)	(0.030)
N	1,551	831	1,276	628	1,551	831	1,276	628	1,550	830	1,275	627
Mean value (incl. HH)	0.167	0.171	0.170	0.171	1264.487	1174.655	1364.989	1271.818	0.107	0.100	0.101	0.095
S.d. (incl. HH)	0.373	0.377	0.376	0.377	31493.721	30331.384	33012.813	31858.207	0.174	0.170	0.170	0.166
Covariates	1	1	1	1	1	1	1	1	1	1	1	1
Bandwidth	10	5	10	5	10	5	10	5	10	5	10	5

Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Figure A.1: Balance of baseline characteristics at the age 60 cutoff by bandwidth



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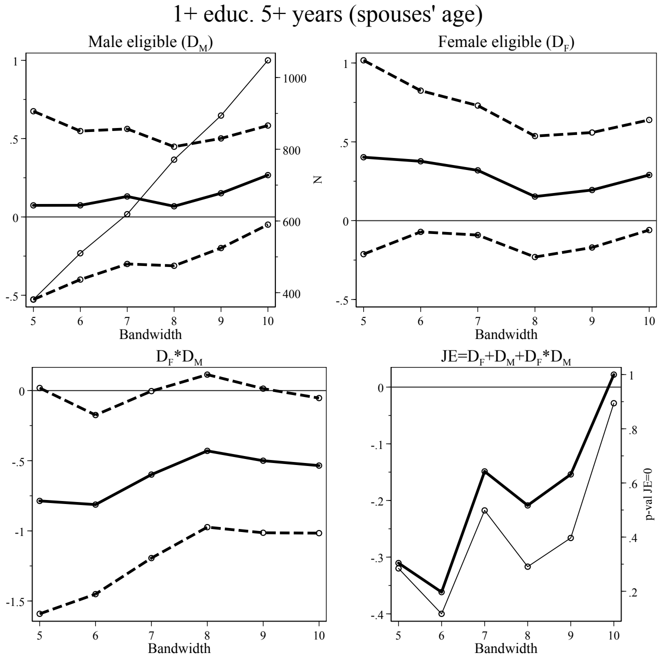

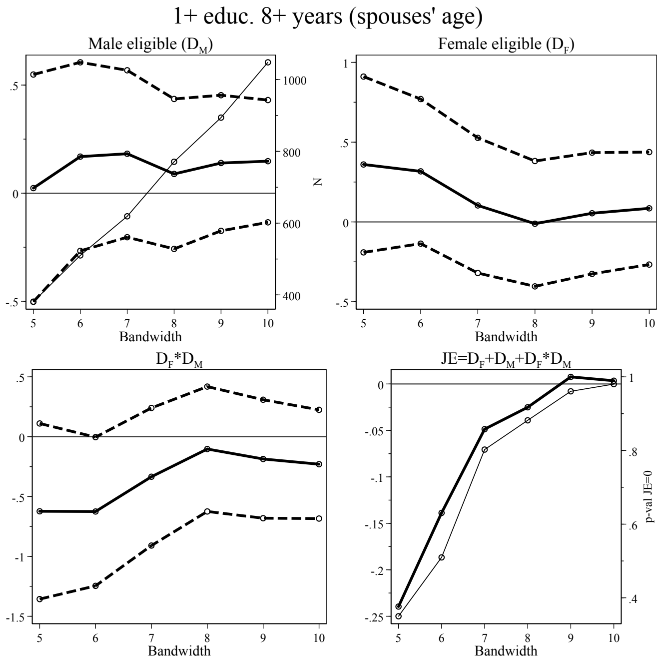

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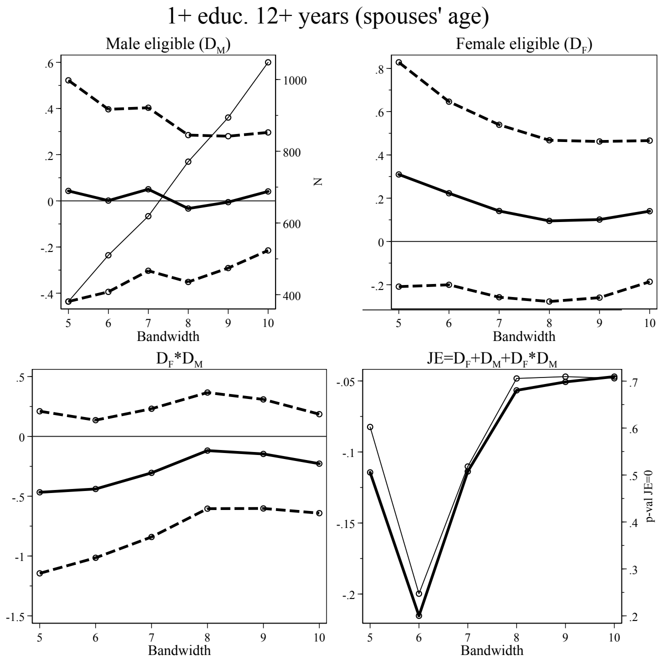

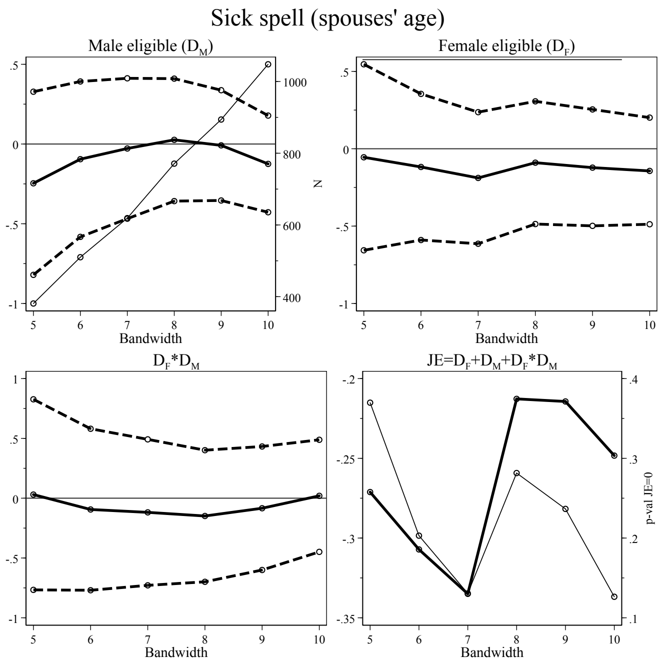

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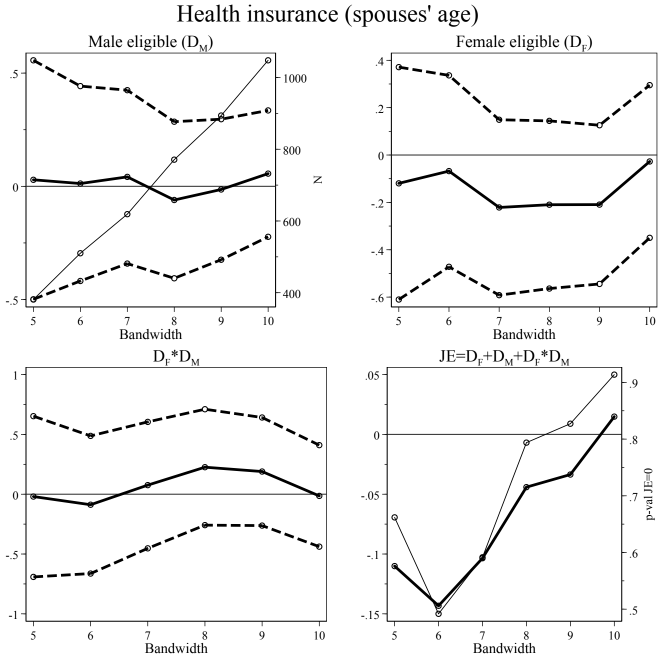

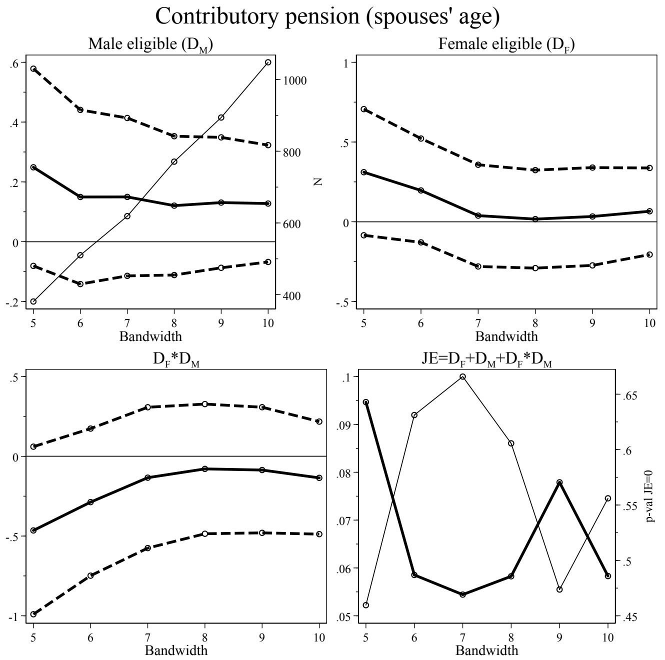

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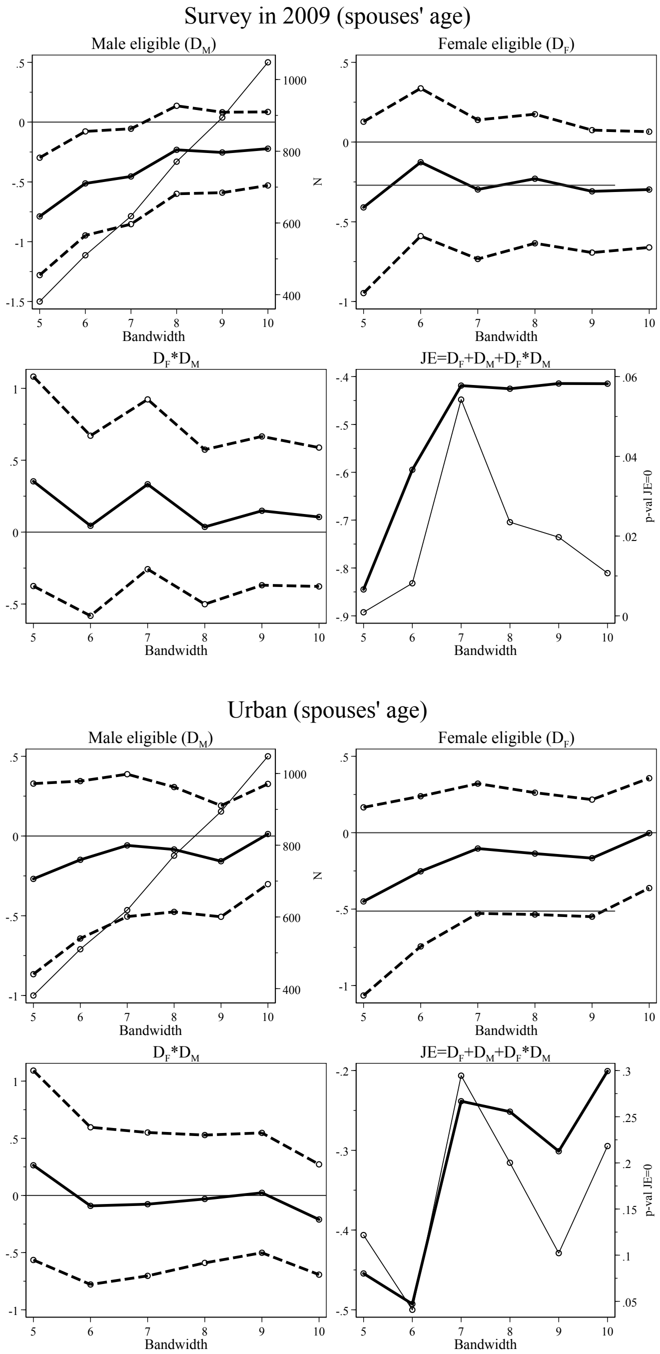
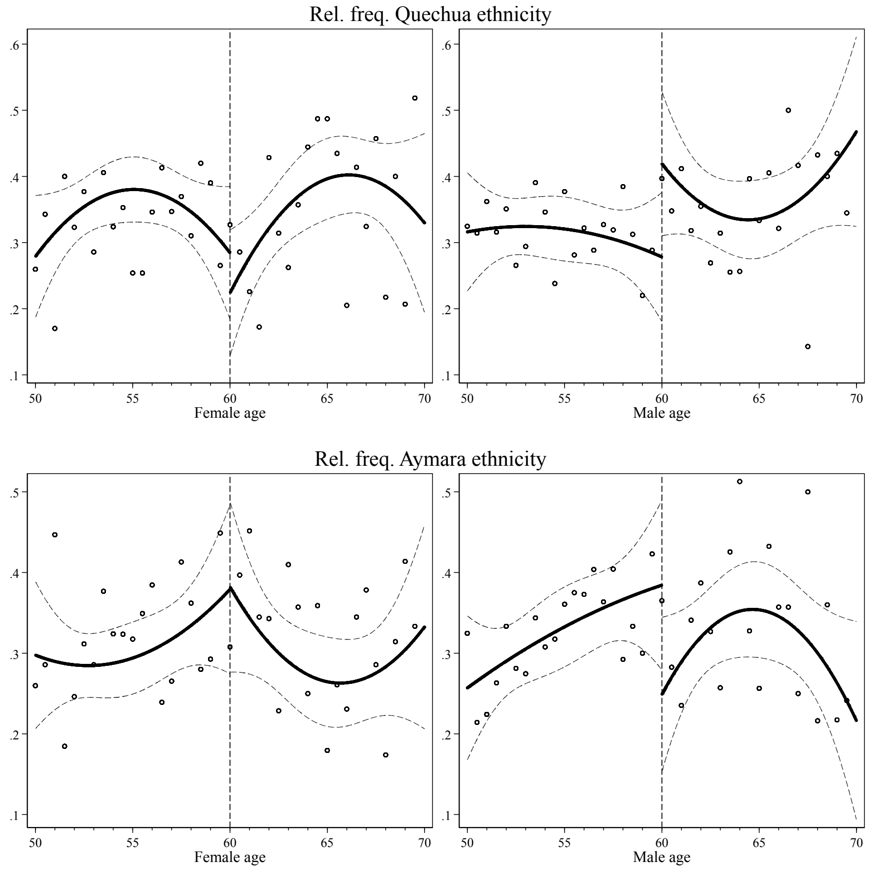


Figure A.2: Balance of baseline characteristics at the age 60 cutoff



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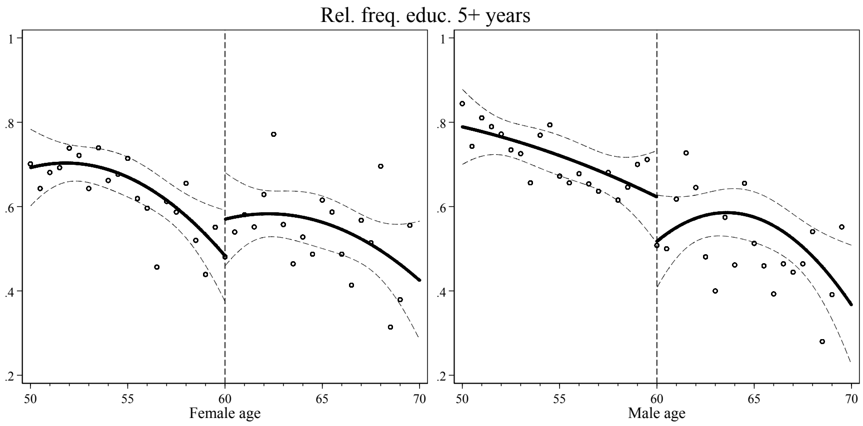

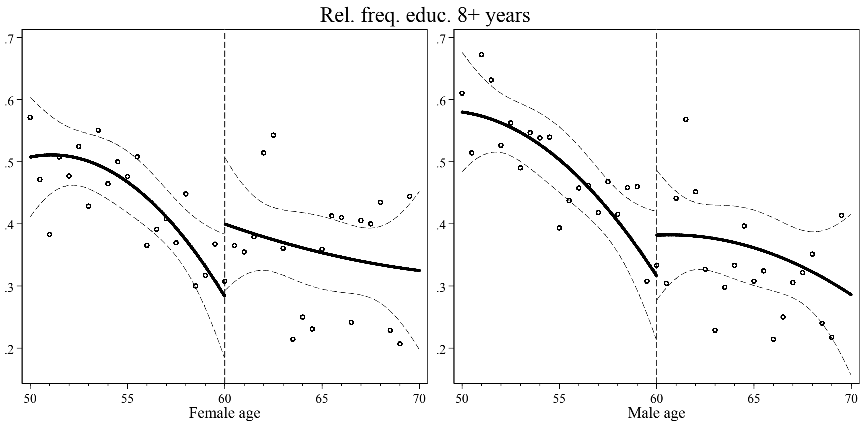

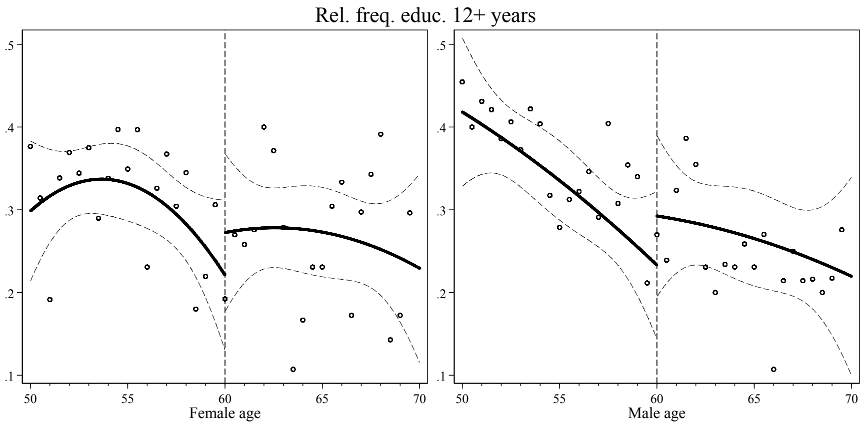

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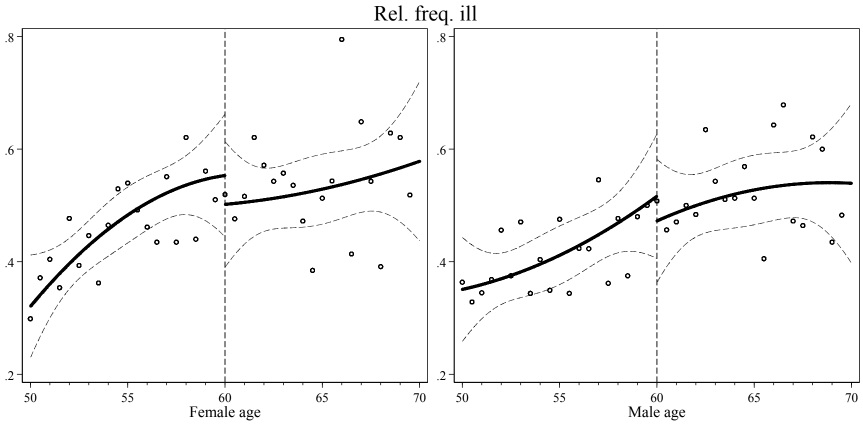

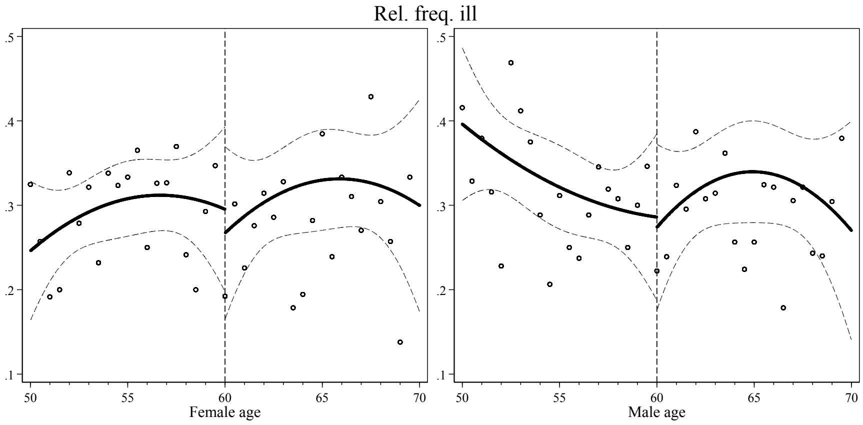

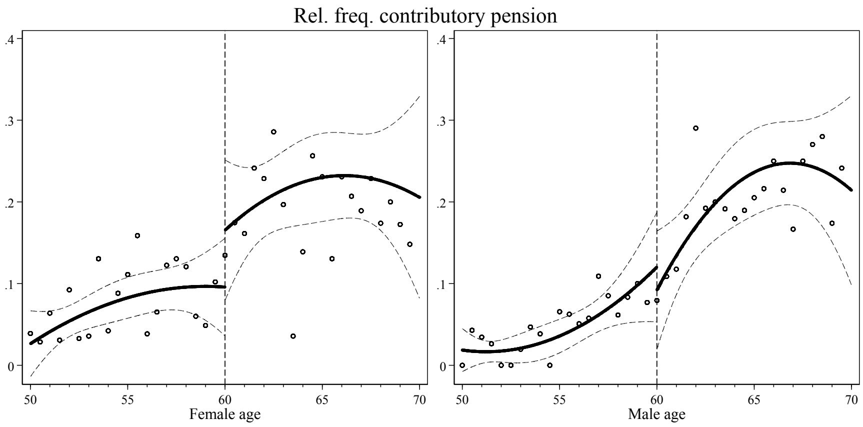

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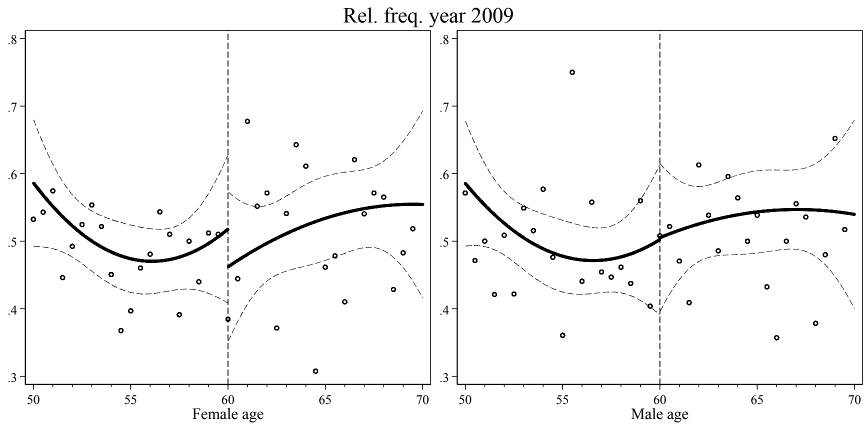


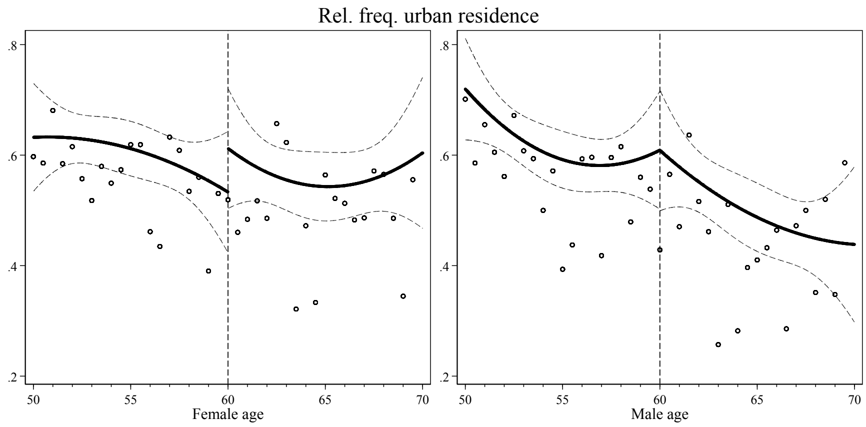
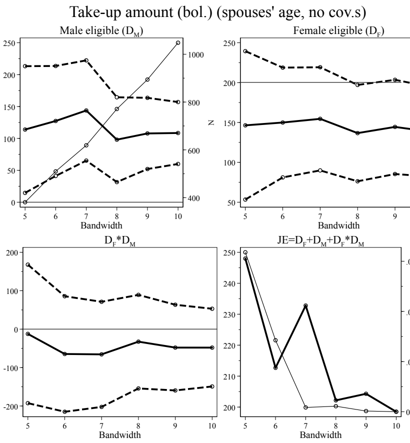
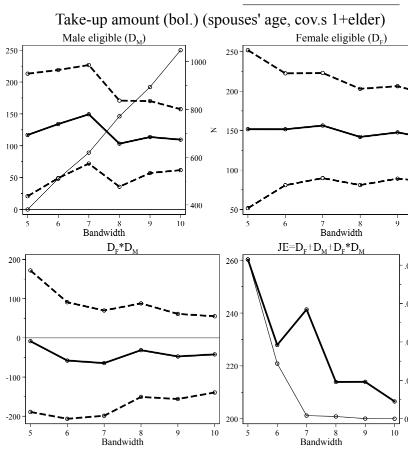


Figure A.3: Robustness of eligibility effect to varying the bandwidth

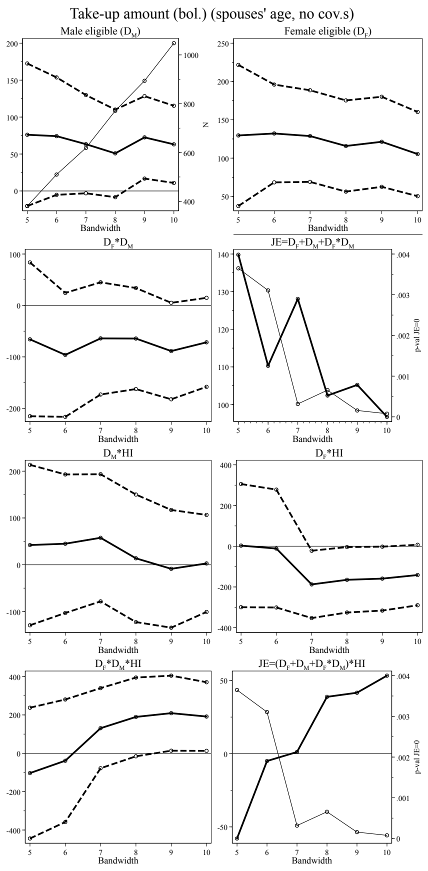
(a)



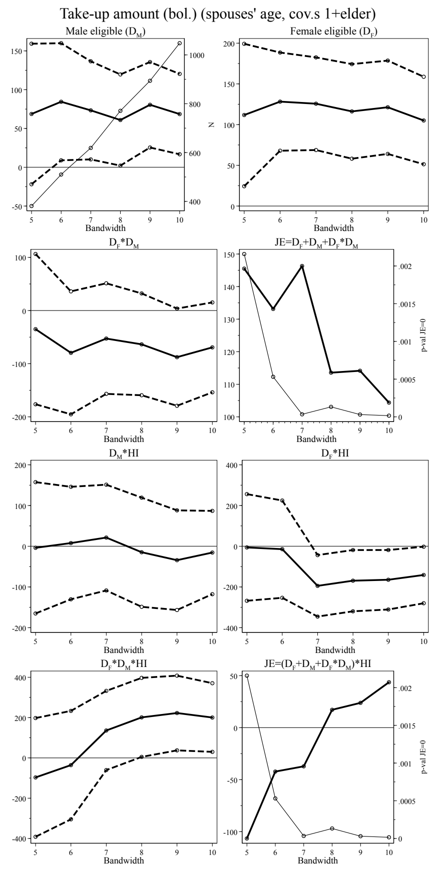
(b)



(c)



(d)



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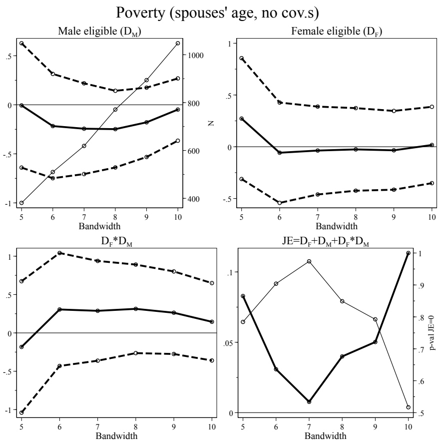
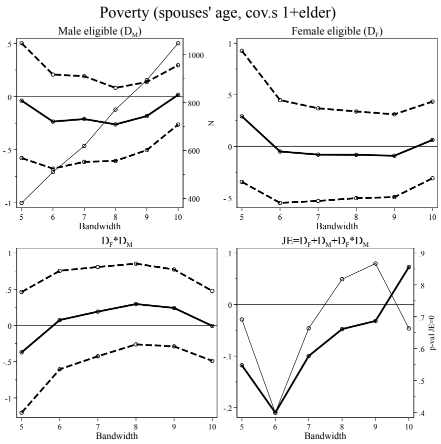


(a)



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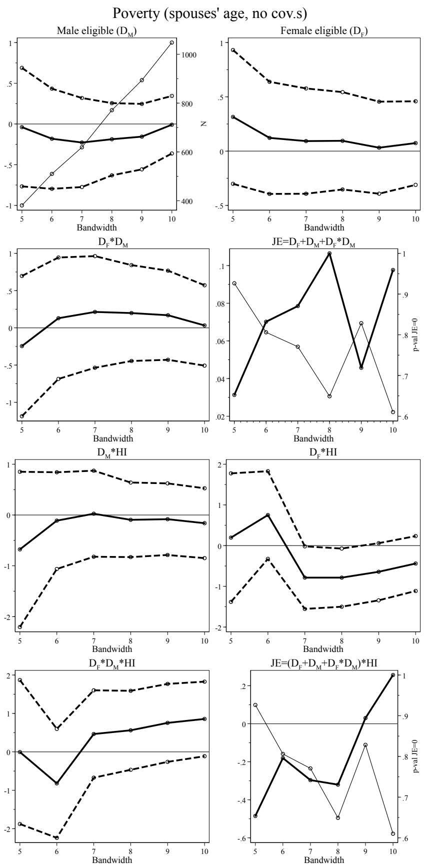

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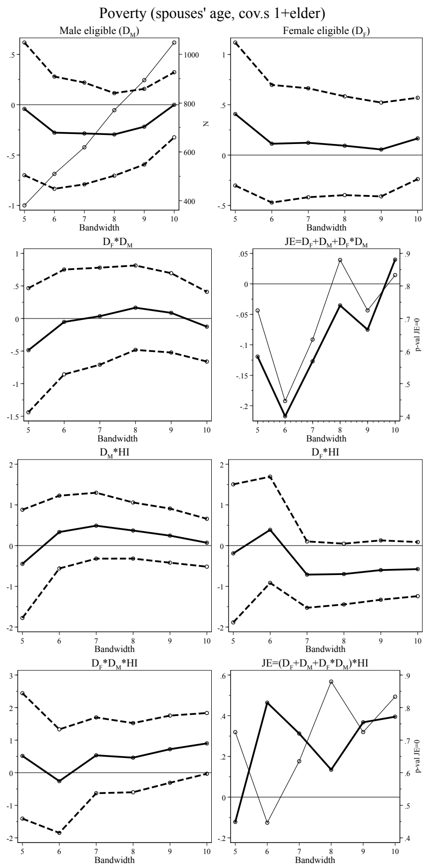

(d)

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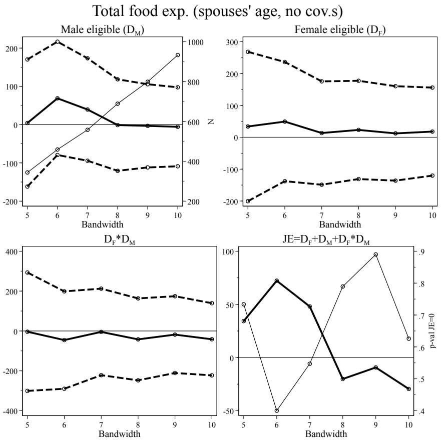
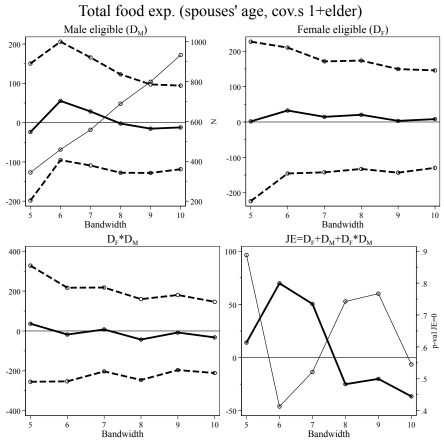
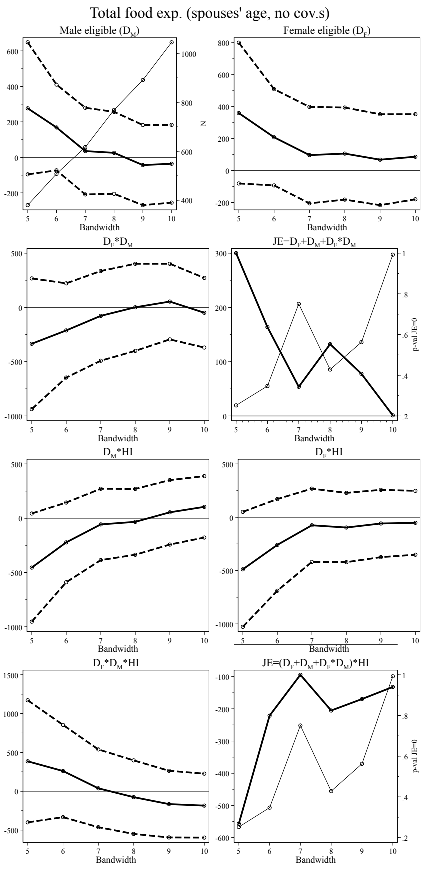


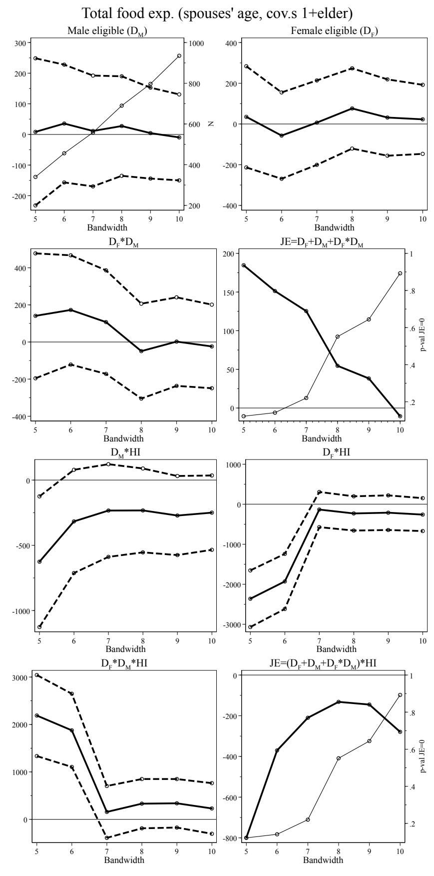
(a)



(b)



(c)



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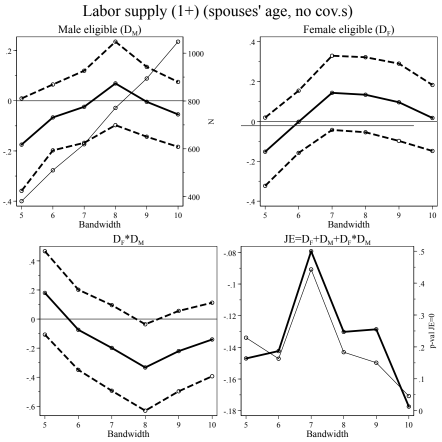

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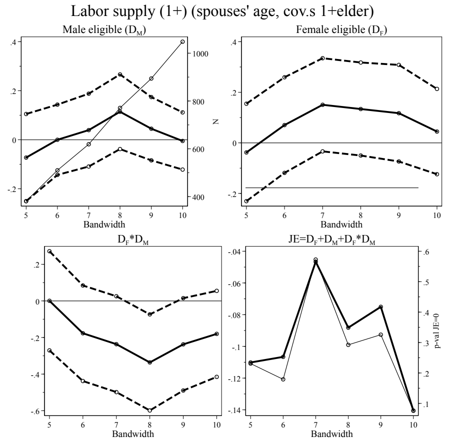

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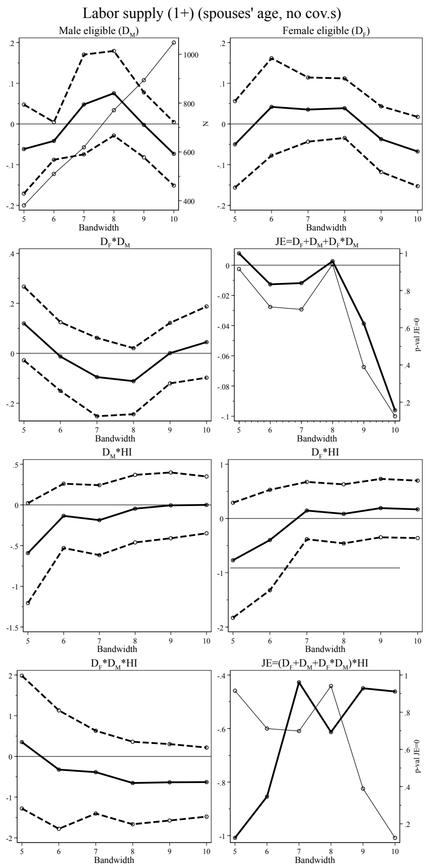

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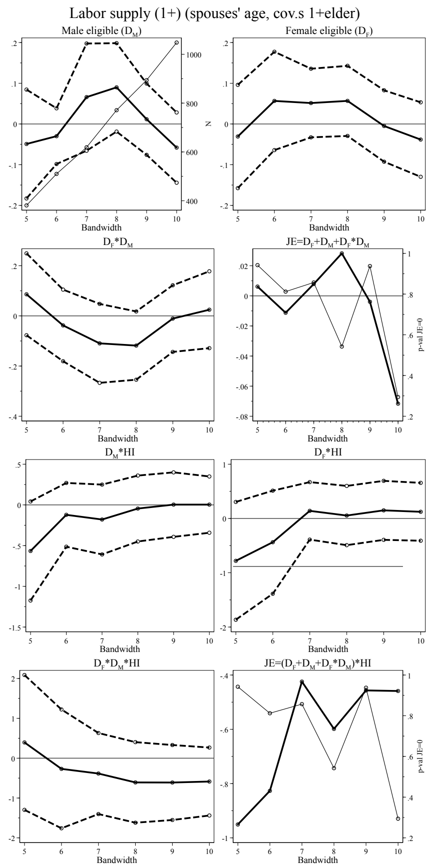

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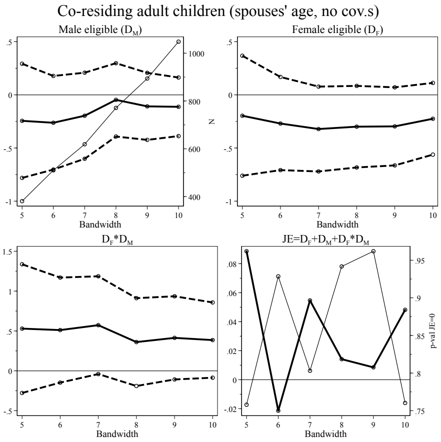

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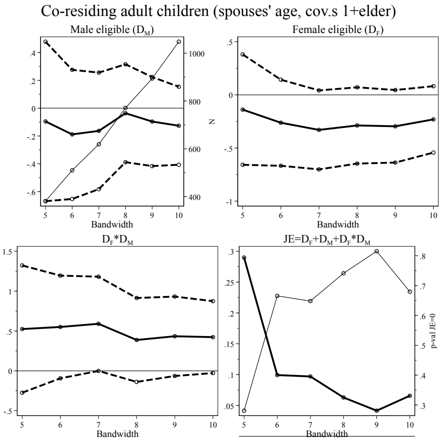

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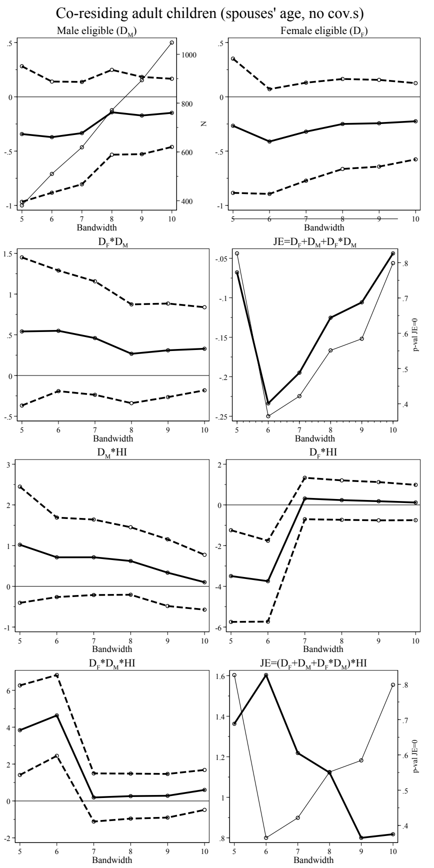

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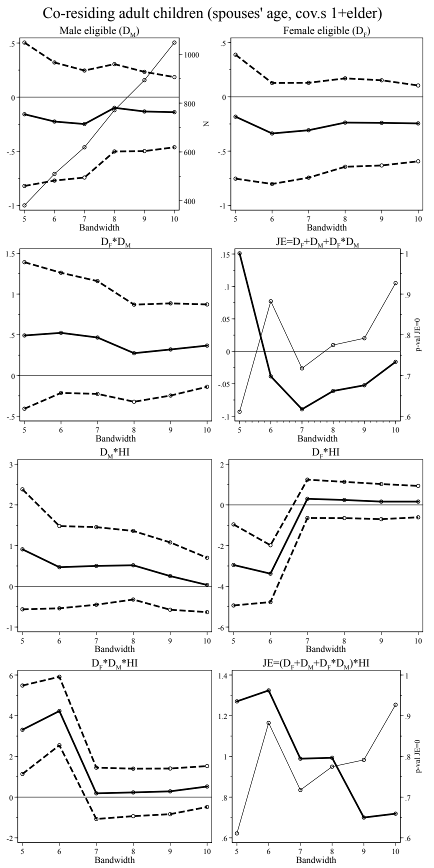

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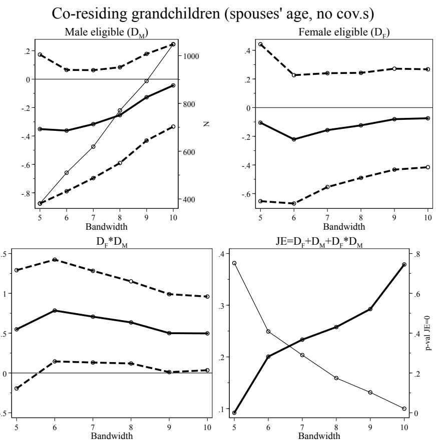

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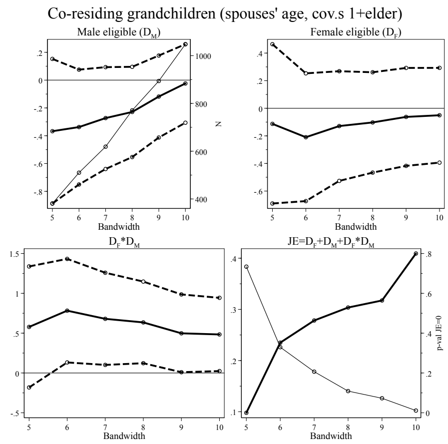

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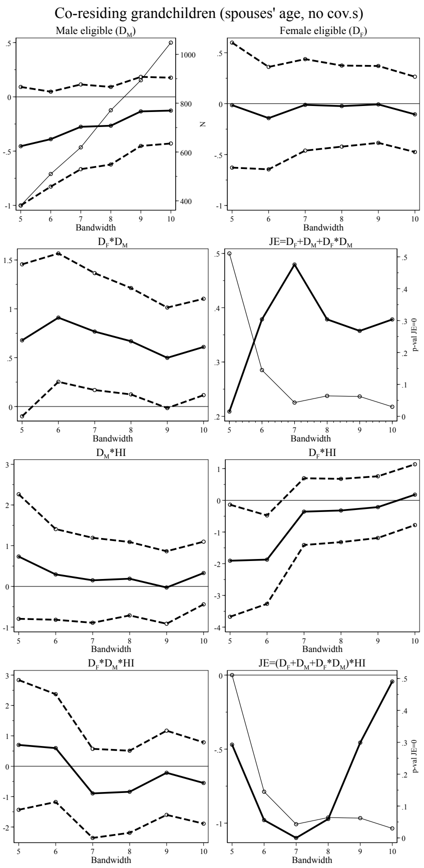

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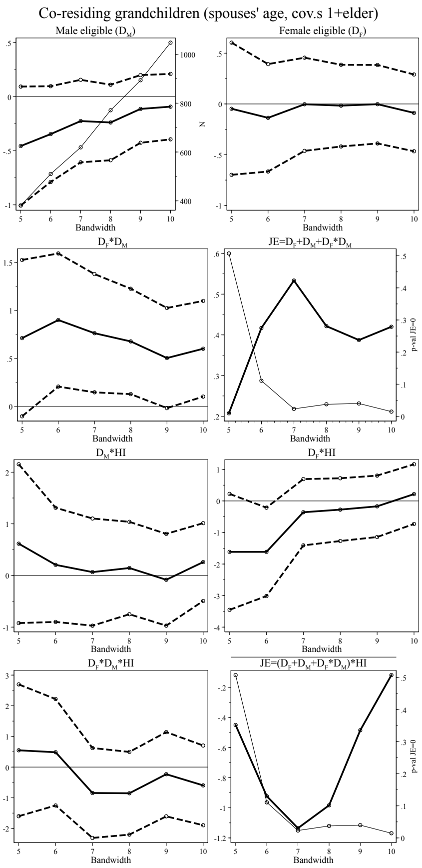
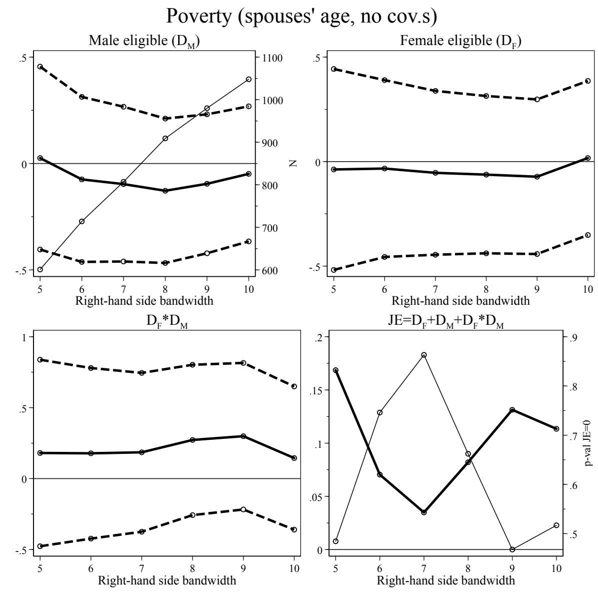
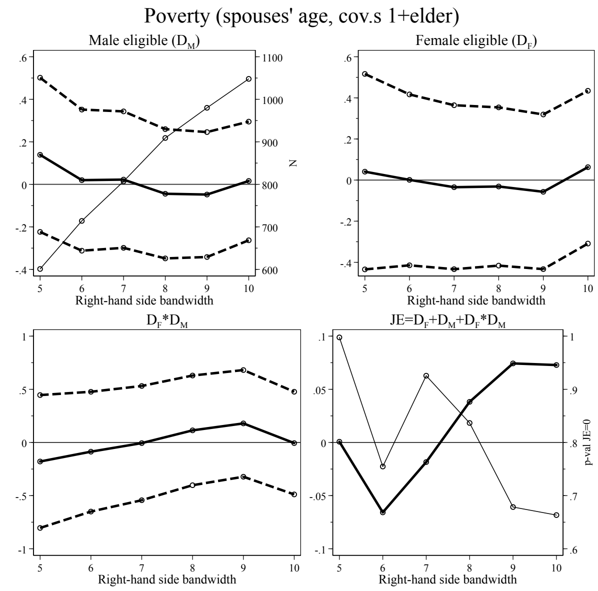


(d)

Figure A.4: Robustness of eligibility effect to varying the right-hand side bandwidth





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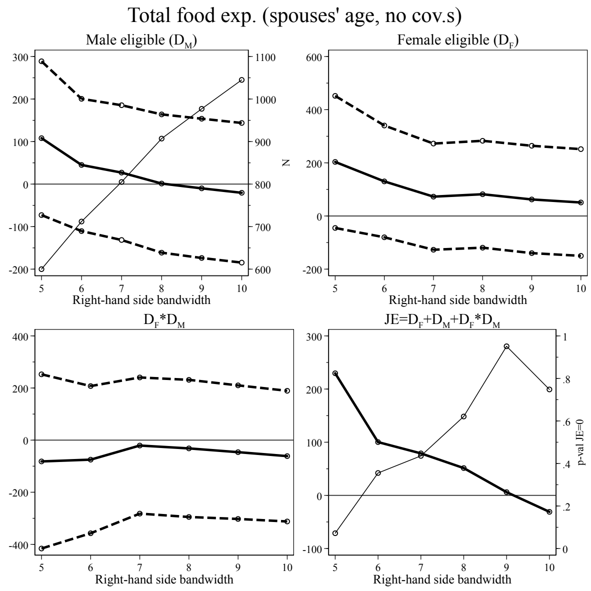

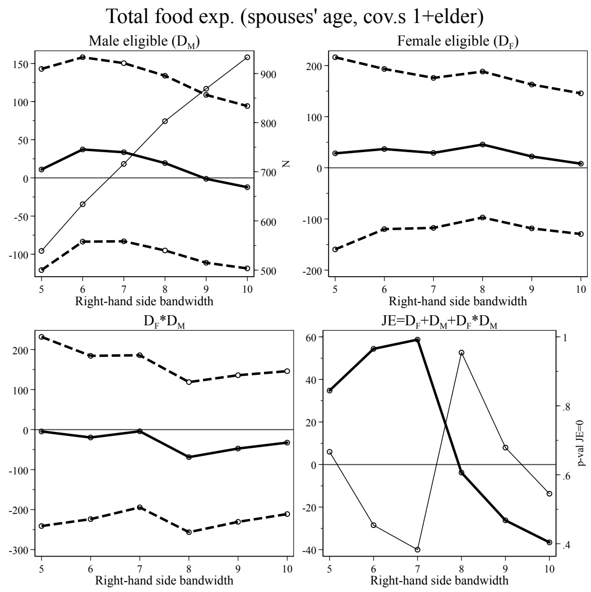

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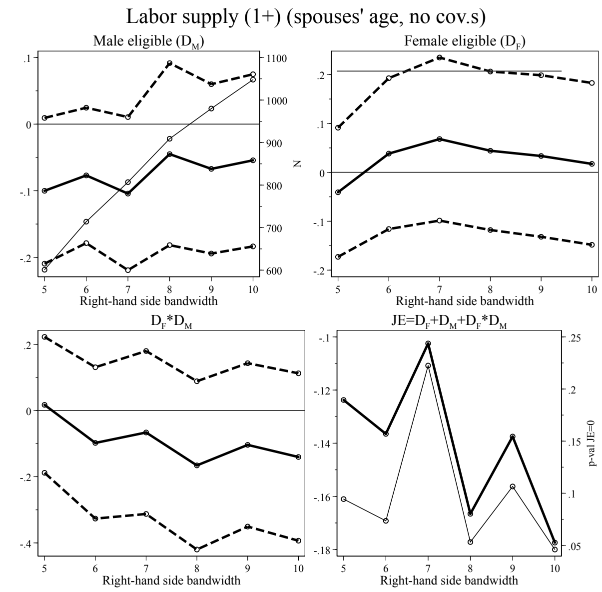

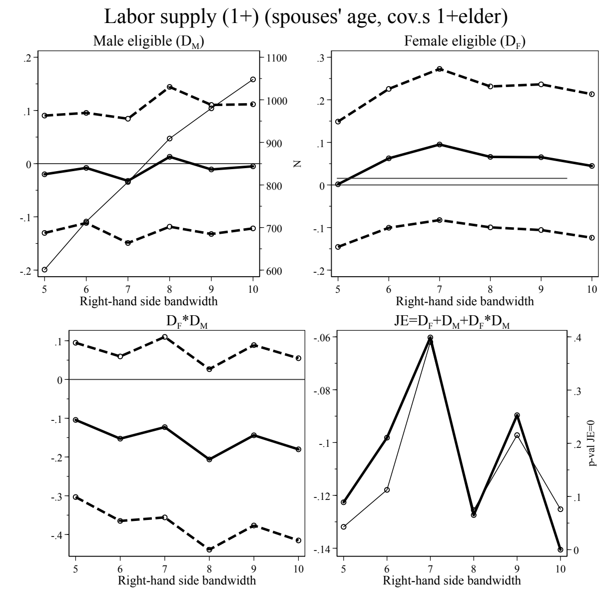

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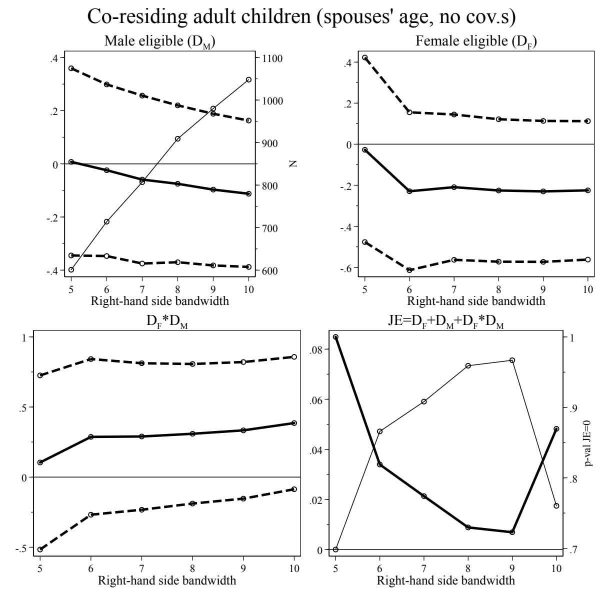

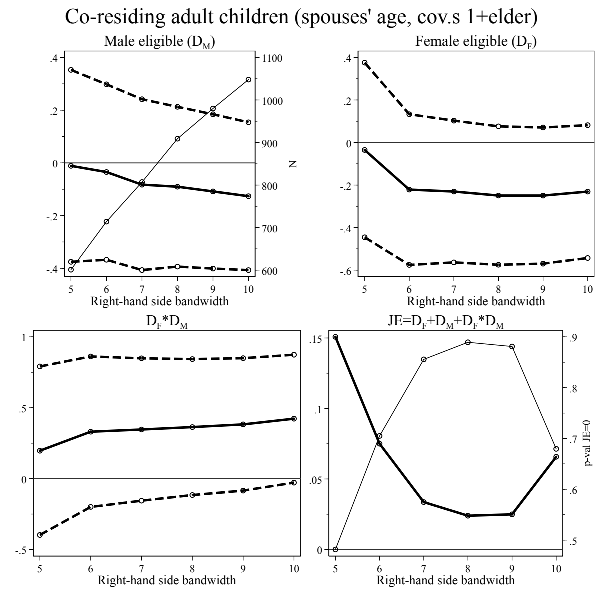

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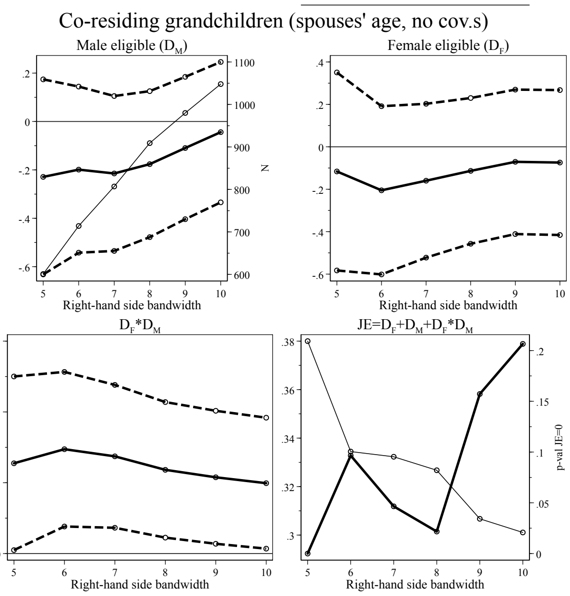

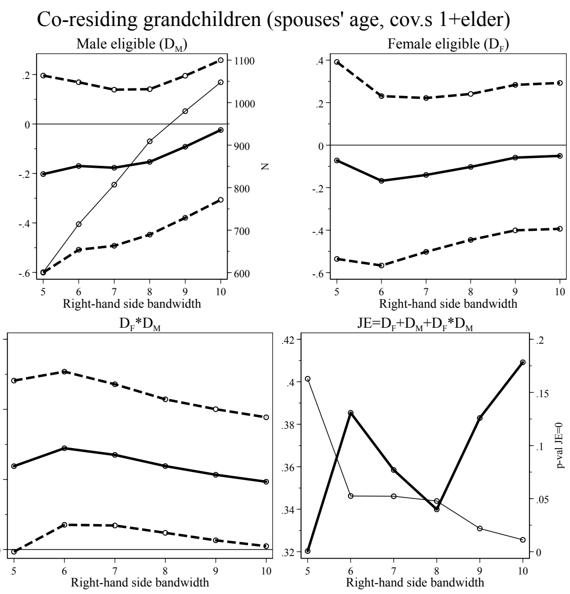

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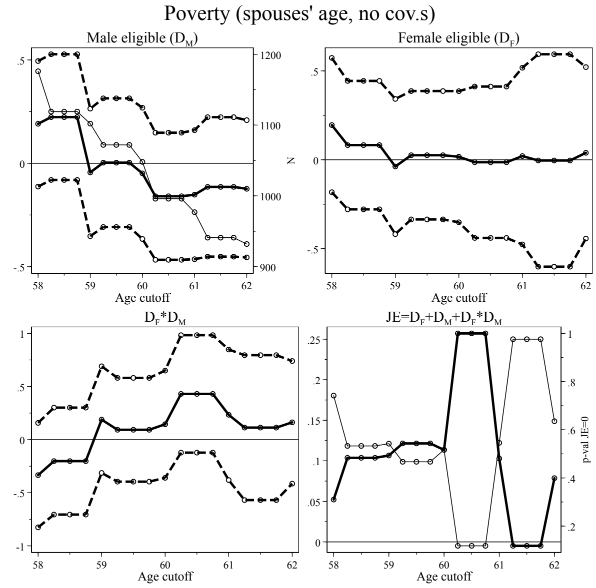
Figure A.5: Ant(post)icipation effects of pension eligibility



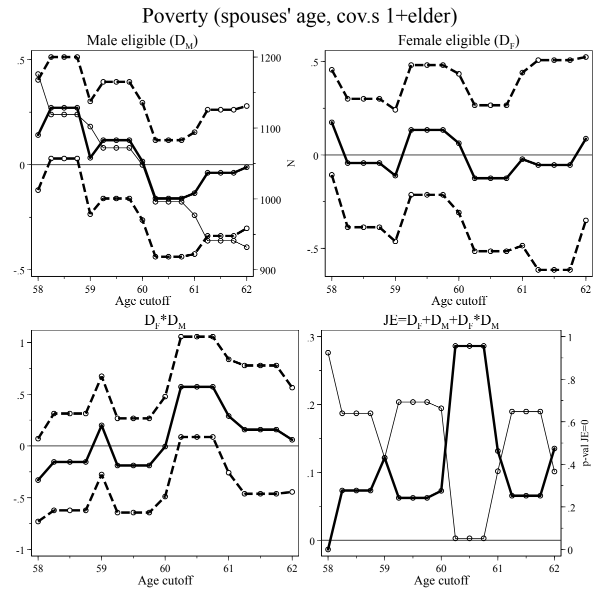


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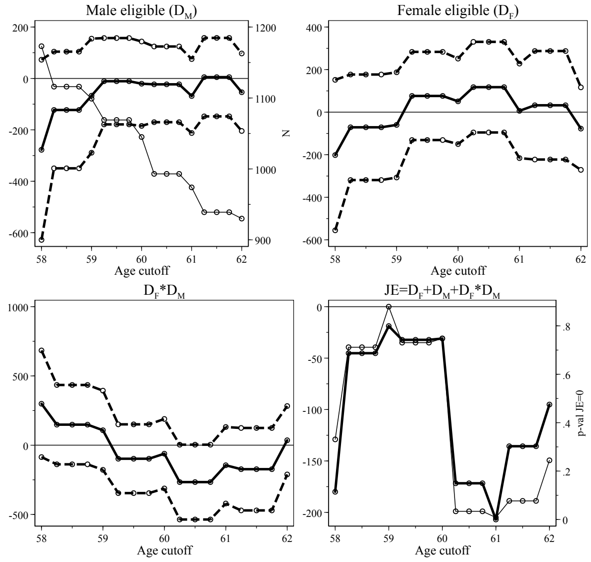
Figure A.5 continued from the previous page...

Total food exp. (spouses' age, no cov.s)



Total food exp. (spouses' age, cov.s 1+elder)

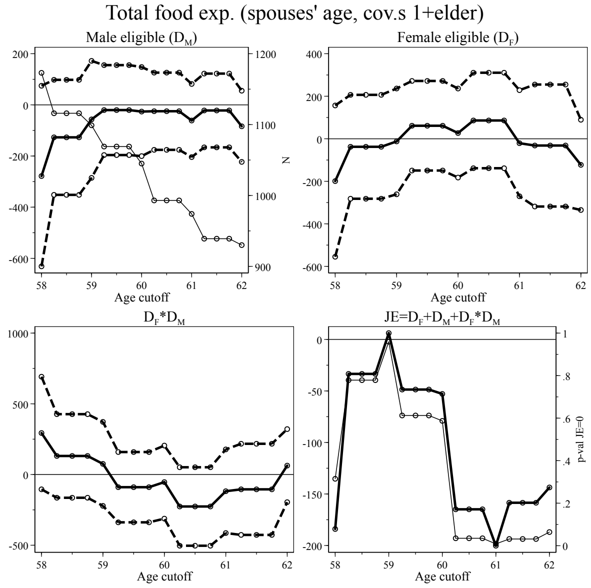


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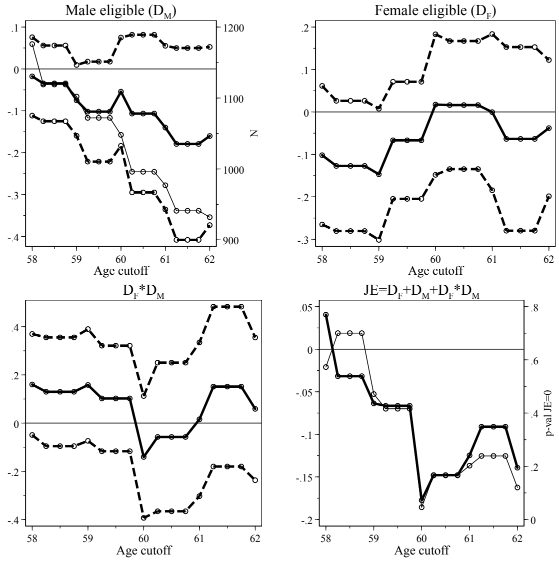
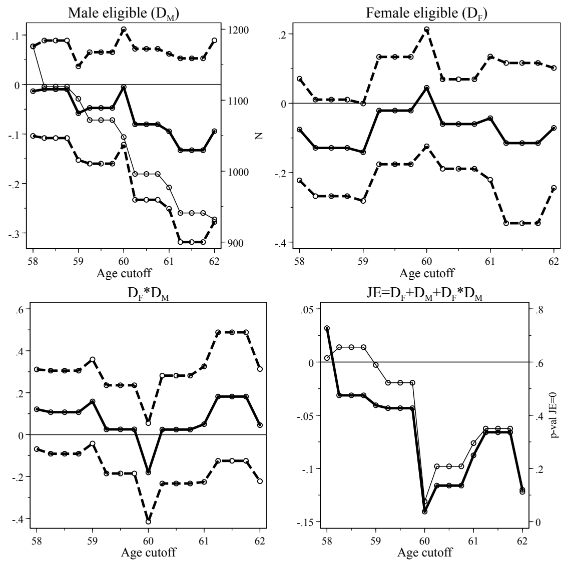
Figure A.5 continued from the previous page...

Labor supply (1+) (spouses' age, no cov.s)



Labor supply (1+) (spouses' age, cov.s 1+elder)



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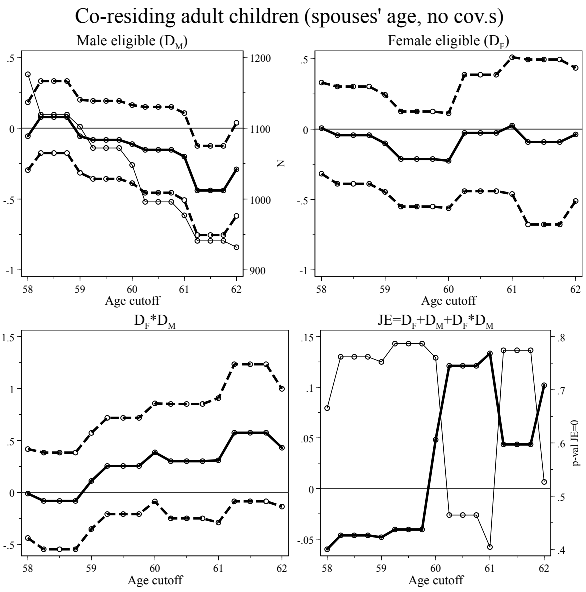

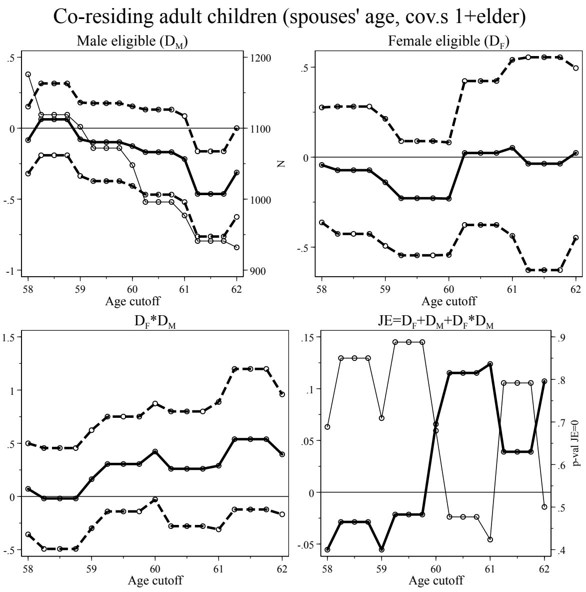

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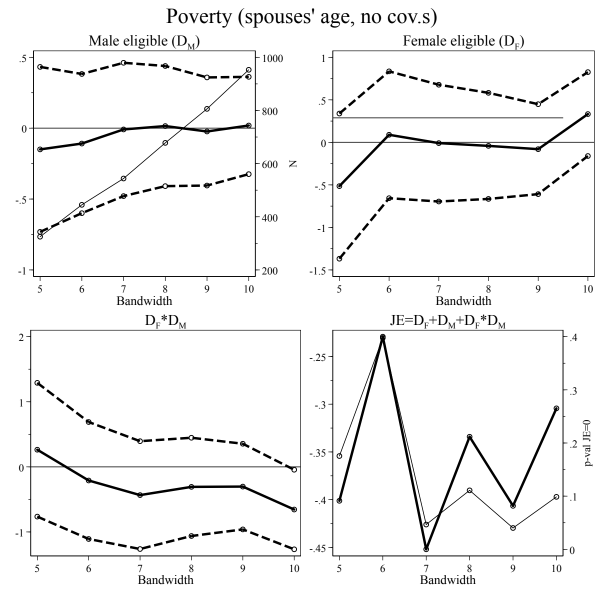
Figure A.6: Placebo effect of pension eligibility



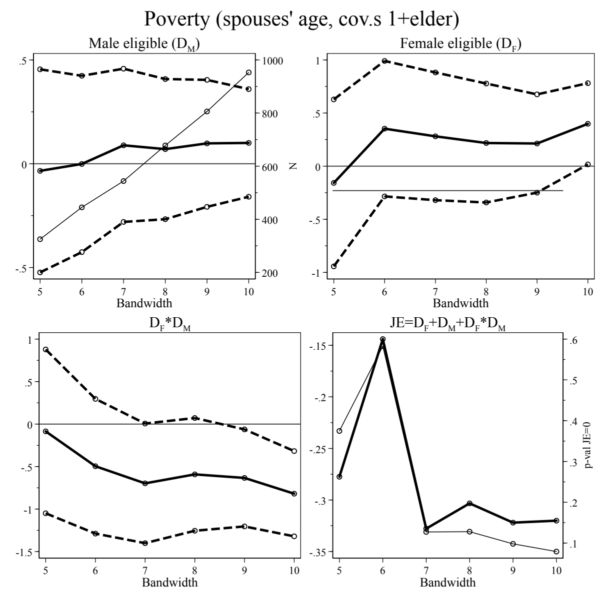


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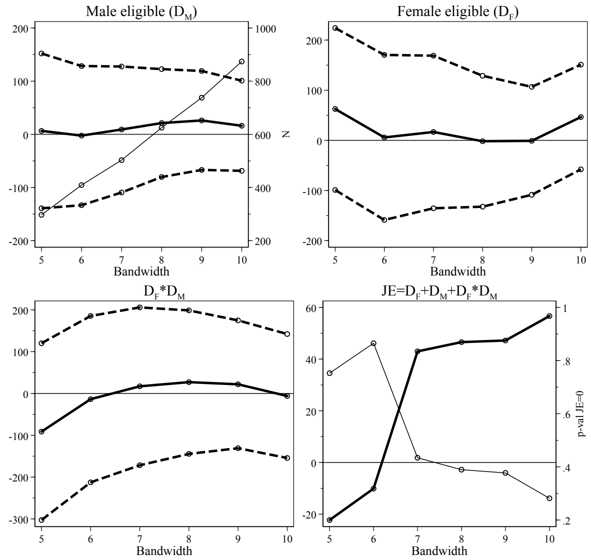
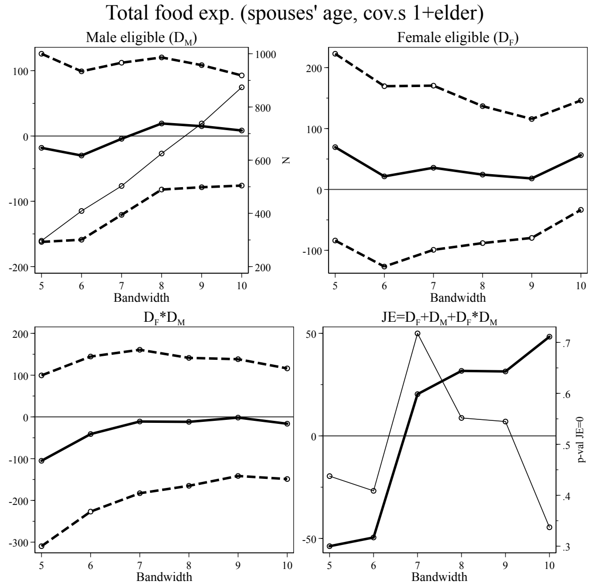
Figure A.6 continued from the previous page...

Total food exp. (spouses' age, no cov.s)



Total food exp. (spouses' age, cov.s 1+elder)



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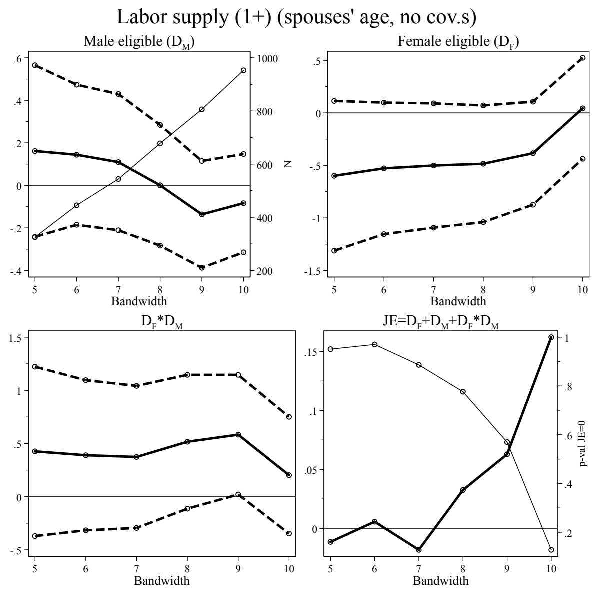

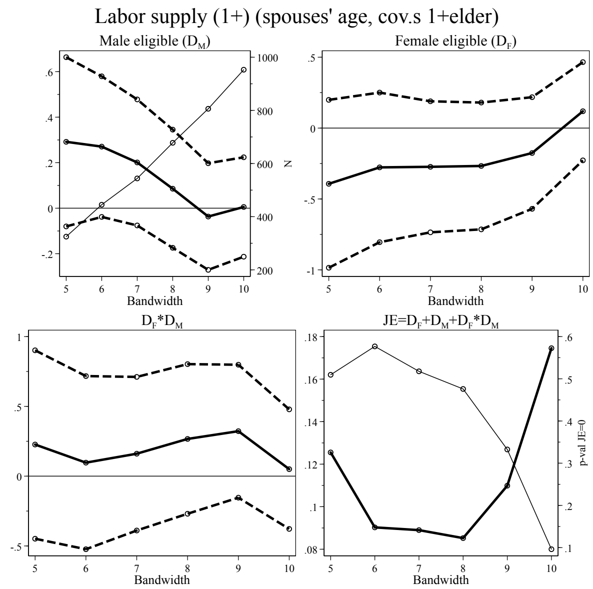

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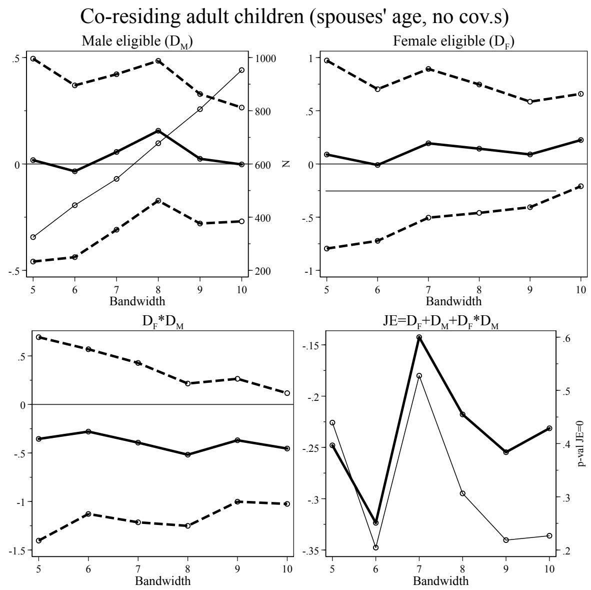

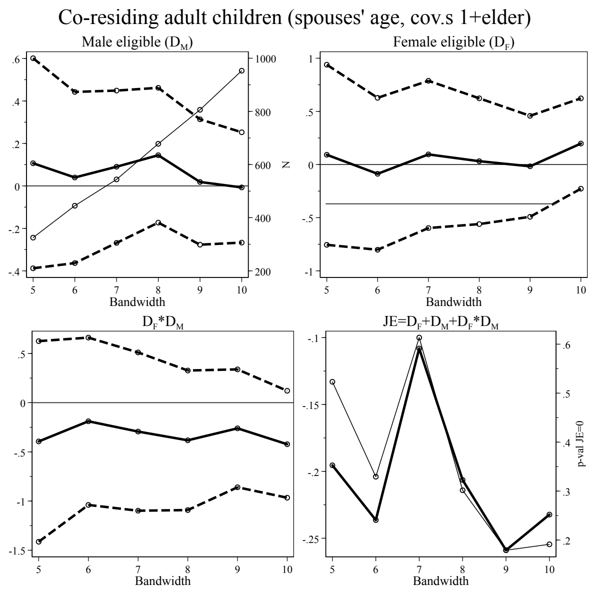

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